Essays on international finance and empirical asset pricing

Maletic, M.

Publication date:
2020

Document Version
Publisher's PDF, also known as Version of record

ESSAYS ON INTERNATIONAL FINANCE AND EMPIRICAL ASSET PRICING

MATJAZ MALETIC

Essays on International Finance and Empirical Asset Pricing

Proefschrift

Proefschrift ter verkrijging van de graad van doctor aan Tilburg University op gezag van prof. dr. K. Sijtsma, in het openbaar te verdedigen ten overstaan van een door het college voor promoties aangewezen commissie in de Portrettenzaal van de Universiteit op woensdag 15 januari 2020 om 10.00 uur door

Matjaz Maletic

geboren op 12 april 1985 te Ljubljana, Slovenië.

Promotor:	Prof. dr. B. Melenberg
Copromotor:	Dr. L.T.M. Baele

Promotiecommissie:	Prof. dr. J.J.A.G. Driessen
	Prof. dr. F.C.J.M. de Jong
	Dr. P.C. de Goeij
	Dr. A. Diez de los Rios

Acknowledgments

It has been a wonderful journey. Foremost important is my great indebtedness to my supervisors, Bertrand and Lieven. I would like to thank you for your help and support. Especially, I would like to thank you for your persistence and for your understanding.

I would like to thank Luc and Juan Carlos for giving me an opportunity to enroll in the PhD programme at Tilburg University. I would like to thank Joost, Frank, and Peter for becoming part of my PhD committee and helping me during the final stages. The saying states that behind every successful man stands a woman, so I would like to thank Marie-Cecile, Loes, and Helma for their administrative support.

I would like to thank Antonio and Narayan for their guidance, for helping me, inviting me, and hosting me in Ottawa. It was an excellent experience. I am looking forward to our future collaboration!

I would like to thank Igor for his supervision at the Faculty of Economics, and for helping me throughout the PhD. I would like to thank Arjana for recruiting me at the Bank of Slovenia and providing a good policy-oriented research environment.

Finally, I would like to thank my family, friends, and colleagues. My nephew Simon, to whom I am a godfather, was born when I enrolled in a research master programme and left Ljubljana. Today, he can walk, talk, and almost already take care of himself. Similarly, during my PhD, I have learned how to walk (learned the methodology), talk (more precisely write), and hopefully, demonstrated my ability to pursue independent research.

The list of people outside of my family to whom I am grateful and who supported me is too long to be reported exhaustively. By the time I graduated, many of my friends got married and started a family. I wish you and your children only the best. I am especially grateful to Bojan, Antica, Tomaz, Polona, Damjan, Gordana, Darja, Emerson, Rachel, Andreas, Felipe, Ziga, Jure, Borut, Timotej, Riccardo, Milan, Matic, Crt, Robert, Miha, Burak, Joren and Jan.

Instead of looking at the graduation as the end, I like to think of it as the beginning of a new journey that lies in front of me. It helped me to improve my methodological and writing skills. However, it also pointed to some of my weaknesses.

During my PhD, I noticed the importance of expectations. By working hard, not being afraid to fail, and having realistic expectations, you determine your success. You can learn from your past mistakes, become more balanced, and build your future resilience.

Respectfully,

Matjaz Maletic, Tilburg, January 2020.

"All models are wrong, but some are useful."

GEORGE BOX

Contents

Introduction

This thesis consists of three chapters. The first two chapters analyze the different channels by which changes to the Chinese economy affect the dynamics of the US (and German, in Chapter 1) term structures. The third chapter contributes to the active cross-sectional asset pricing literature by analyzing how R&D investments and past returns interact in explaining future returns.

The aim of the first chapter, entitled "*A Chinese slowdown and the US and German yield curves*", is to quantify the impact of the Chinese economy on the US and German term structures of interest rates through the lenses of traditional asset pricing models.

Given the interconnectedness of China and the rest of developed economies, a Chinese slowdown might spill over to the US and German economies, and hence also to their term structures of interest rates, in two ways. First, lower Chinese growth might signal lower expectations about future growth and inflation. Second, bad news about the Chinese economy might increase the uncertainty about future developed market growth and inflation. Consequently, market participants might revise their expectations of future monetary policy actions in light of this new information. In an environment where the short rate is at the effective lower bound, this implies that the Central Banks will hold interest rates "lower for longer." Correspondingly, a lower Chinese growth leads to a lower term premium, the compensation for bearing the duration risk.

I estimate an affine term structure model to decompose the 5y nominal yield in (1) an expected future 5y nominal short rate, "the expectations channel," and (2) the 5y term (risk) premium. Empirically, I represent a Chinese slowdown with a drop in the Chinese leading indicator.

A drop in the Chinese leading indicator decreases the 5y Treasury and Bund yields by decreasing the 5y term premia. My empirical findings are consistent with the argument that a Chinese slowdown is a signal for lower long term nominal interest rates and mainly alters risks that future growth and inflation will be lower than expected. In the post financial crisis environment with low growth and inflation, and monetary policy constrained by the effective lower bound, investors became very sensitive to a deterioration of the outlook about the Chinese economy. They are willing to accept lower compensation for holding nominal long-term bonds instead of short-term securities.

The second chapter, entitled "*Chinese foreign reserves and the US yield curve*", focuses on the bilateral relationship between the US and China. In particular, I am investigating how the accumulation of the Chinese foreign reserves is affecting the US yield curve through the lenses of modern portfolio-balance models.

China is managing its exchange rate against the US Dollar. The Renminbi was pegged to the US Dollar since 1994. In 2005, China moved towards a managed peg. The Renminbi, however, preserved a tight link to the US Dollar. The combination of a managed exchange rate and record-high growth rates resulted in a surge in the Chinese foreign exchange reserves. By June 2014, the Chinese foreign reserves increased to 4 trillion US Dollars.

After the financial crisis, however, the Chinese foreign reserves were growing at lower and even negative yearly rates, while the Renminbi appreciated. The strong Renminbi put an additional anchor on economic growth, which was already slowing down. A substantial appreciation of the Renminbi in 2014 due to a strong US Dollar and economic slowdown in China were building market consensus that the Renminbi was overvalued. In July 2015, the PBOC moved closer towards the market determination of the Renminbi. Market participants interpreted the regime change as the beginning of a sizeable depreciation.

I find that when the Chinese official sector rebalances away from the US Dollar, it lowers the 5y Treasury yield and the 5y Treasury term premium. Such rebalancing can be a result of (unexpected) significant depreciation of the Renminbi against the US Dollar and increased (unobserved) uncertainty that the future growth of the Chinese economy will be lower than expected. In an environment where the monetary policy is constrained by the effective lower bound, lower Chinese foreign reserves increase the value of the 5y US Treasuries and decrease the 5y Treasury term premium.

The third chapter, entitled "*R&D Investments, Past Returns, and the Cross-Section of Stock Returns*", investigates how R&D investments and past returns interact in explaining future returns. Existing empirical literature gathered evidence that firms with high R&D-to-market value are rewarded with higher future returns. Firms with a higher level of R&D expenditures, however, are not rewarded with higher future returns unless they have experienced poor past performance.

I contribute to the literature by estimating the cross-sectional regressions which show that the level of R&D interacts differently than changes of R&D in explaining future stock returns. Firms, which are reluctant to cut the level of R&D expenditures despite the poor past performance, are rewarded with higher subsequent returns. On the other hand, the good track record in the past price performance is providing a signal for higher future returns when managers decide to increase the R&D expenditures. Only firms, which have demonstrated their ability to make good investment decisions, and therefore exerted positive price performance over the last year, are rewarded with higher future returns.

Chapter 1: A Chinese slowdown and the US and German yield curves

Matjaz Maletic[1,2]

This version: 24th of October 2019

Abstract

To measure the global spillovers of a Chinese slowdown on the 5y nominal interest rates in the US/Germany, I model the US/German yield curves jointly in the post financial crisis sample, including the Chinese leading indicator as a new factor. I use an affine term structure model and decompose changes in the 5y nominal interest rates into (1) changes in the 5y expected future nominal short rate, and (2) the 5y term premium. A drop in the Chinese leading indicator decreased the 5y Treasury yield and the compensation for bearing the duration risk (the 5y Treasury term premium). In Germany, the lower Chinese leading indicator moderately increased the 5y Bund yield by increasing the term premium attached to the 5y German Bunds. However, as such increases of the term premium could be driven by recessions I re-estimate a single country affine term structure model for Germany in the post sovereign debt crisis sample. Like in the US, I now find that in Germany, a lower Chinese leading indicator decreased the 5y Bund yield and its term premium.

1. Introduction

The last decades have witnessed tremendous growth of the Chinese economy. In 2017, China accounted for 15 percent of global GDP, compared to only 3 percent in 1999. While the growth of the Chinese economy continues to outshine that of its global peers, the growth has dropped from double digits before the crisis to 7–8 percent after the crisis.

Existing work has investigated the impact of (changes in) Chinese growth on, amongst others, the global/US/EU growth and inflation dynamics, unequivocally finding the effects to

---

[1] Tilburg University, Finance Department, e-mail: maletic.matjaz@gmail.com

[2] I would like to thank Joost Driessen, Frank de Jong, Narayan Bulusu, Arjana Brezigar Masten and the Brown Bag seminar and Macro Reading group participants at Tilburg University. I would like to thank Igor Loncarski, and especially my supervisors Bertrand Melenberg and Lieven Baele for their comments. I would like to thank Antonio Diez de Los Rios for his support, comments, and sharing the Matlab codes.

be large[3,4]. To the best of my knowledge, this paper is the first to quantify the effects of a Chinese slowdown on the US and German yield curves. This is my main contribution.

I hypothesize that a Chinese slowdown can affect the US/German yield curves through two channels, in several (possibly opposing) ways.

First, changes in Chinese growth may affect the future expectations of fundamental drivers of the US/German yield curve, such as inflation and real growth rates in these respective countries. Gauvin and Rebillard (2015) and Metelli and Natoli (2017), for instance, show that a Chinese slowdown has substantial negative effects on the US and euro area (EA) growth and inflation rates[5]. The resulting drop in expected real short-term interest rates and inflation leads to a drop in expectations about future nominal short rates. Following Bauer and Rudebusch (2014), I call the future expected nominal short rate the "signaling channel."

Second, changes in Chinese growth may affect the US and German term premia attached to the nominal bonds. The lower Chinese growth could lower the expectations of the nominal interest rates by decreasing the compensation for bearing the duration risk (the term premium) through lower growth and inflation *risks*. First, deterioration of the economic outlook of the Chinese economy, and its consequences for the outlook of the global economy could imply that at the effective lower bound the Central banks will have to hold rates lower for longer. In such an environment, nominal bonds hedge against the risk of lower growth while other instruments such as risky stocks do not. Second, lower Chinese growth could

---

[3] Cashin, Mohaddes and Raissi (2017) find that a percentage decrease of the Chinese growth lowers the global growth by 23 basis points, while a surge in global financial market volatility decreases the global growth by 29 basis points.

[4] ECB (2017) estimates that if the Chinese GDP growth decreases by 3 percentage points cumulatively over three years commodity prices decrease by 6 percent over three years.

[5] Metelli and Natoli (2017) estimate that, without taking into account the Central Bank's responses, a negative shock to Chinese investments, and corresponding reduction in annual output growth equal to 2 percentage points over two consecutive years, decrease the US and EA inflation by 10 basis points in the first, and by 40 basis points in the second year. The shock decreases EA GDP by 30 basis points in the first year and by 20 basis points in the second year. US GDP decreases by 20 basis points in the first year and reverts back to 0 in the second.

increase the risks of lower inflation through, i.e., Chinese lower demand for commodities[6]. The lower inflation increases the real value of fixed dollar payments that bondholders receive. To hedge against the risks of low growth and inflation investors are willing to accept low or even negative compensation for holding nominal bonds rather than short-term securities.

Albeit less likely, the lower Chinese growth could increase the risk premium attached to nominal bonds by increasing the uncertainty about the near-term outlook for the global economy or monetary policy. Such increases, however, are usually associated with recessions. In the euro area, the 5y Bund term premium increased during the sovereign debt crisis. In the US, the 5y Treasury term premium temporarily increased during "the taper tantrum" episode in 2013. Additionally, extremely low Chinese growth could be related to higher risk aversion (U-shaped pricing kernels) and could alter the term premium in a non-linear fashion. Baele et al. (2018) find that the model which accounts for the probability weighting (and loss aversion), namely that the investors attach higher probabilities to extreme events (disasters) explains the equity and the variance premia. Since for a global bond investor turmoil in China could represent a catastrophic event, she could correspondingly overweight such an event and given her loss aversion attach bigger term premium when Chinese growth decreases by a significant amount (i.e. more than 5 percent per year).

Figure 1 shows the development of the 5y Treasury and Bund yields, and of the Chinese leading indicator, in the post financial crisis sample. Actual 5y Bund yield decreased from 2.5 percent in 2009 to –18 basis points in 2017. After the sovereign debt crisis, the ECB initiated the QE programmes which depressed the 5y Bund yield. In 2013, the FED chairman Ben Bernanke signaled a decrease of the QE programmes ("the taper tantrum"). In December 2015, the FED began the hiking cycle. From December 2011 to December 2017 the actual 5y Bund yield decreased from 87 to –18 basis points while the actual 5y Treasury yield increased from 87 to 217 basis points.

[6] Gauvin and Rebillard (2015) notice that in 2011 China accounted for 11 percent of global oil, 41 percent of global copper, and 54 percent of global iron ore consumption.

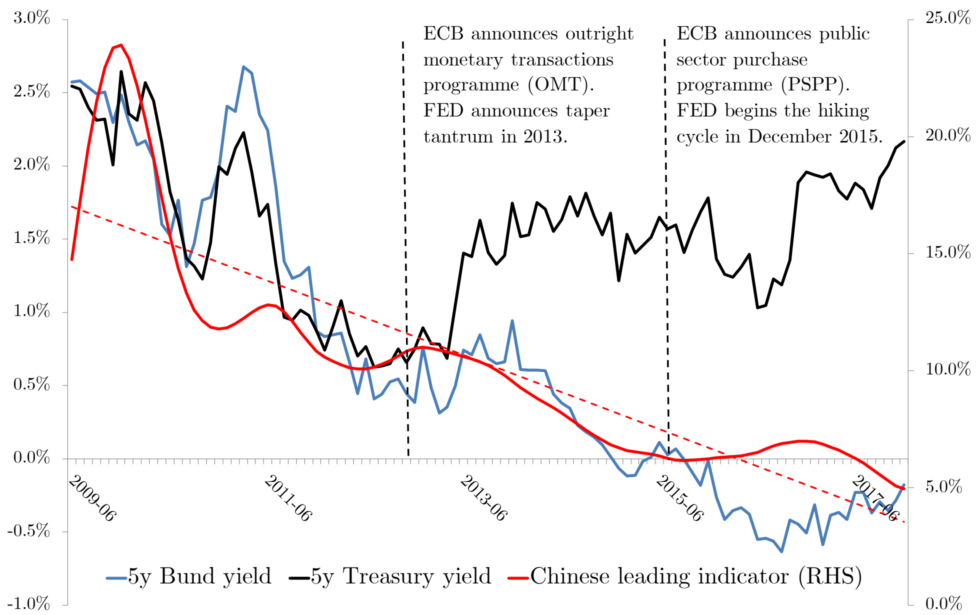


Figure 1: Chinese composite leading indicator (CLI) with trend-restored in twelve-month log differences and the actual 5-year Treasury and Bund yields. The dotted red line depicts a linear trend of the Chinese leading indicator after the crisis. Sample spans from June 2009 to December 2017. Source: BUBA, FED, and the OECD.

My main contribution is to measure the global spillovers of a Chinese slowdown on the US and German 5y nominal yields, the 5y risk-neutral yields, and the 5y term premia. I measure the slowdown with the difference between the GDP growth rates of the domestic economies (the US and Germany) and China. Empirically, I represent the growth rates at a monthly frequency with the leading indicators. To quantify the spillovers of a Chinese slowdown on the US/German 5y yields through the future 5y expected short rates and the 5y term premia, as well as to disentangle both channels, I proceed as follows.

I estimate the joint affine term structure model of the US and German yield curves with the unspanned macroeconomic variables in the post financial crisis sample. With an affine term structure model, I decompose the 5y nominal yields in (1) the expected future 5y nominal short rates, "the signaling channel," and (2) the estimated 5y term premia, "the portfolio balance channel." The alternative name for the first component, the expected future 5y nominal short rate, is the 5y risk-neutral yield. In the model I include, the six principal components extracted jointly from the US and German yield curves and the macroeconomic variables. For *each economy*, the vector-autoregression includes the six principal components, the unemployment rate, core inflation rate, the leading indicator, and the Chinese leading indicator.

In the affine term structure model with the unspanned macroeconomic variables, the macroeconomic variable such as the Chinese leading indicator affects the bond prices only indirectly through the principal components with a lag. In each economy, the US and Germany, I run a vector autoregression of the principal components and the macroeconomic variables. I increase the principal components by the significant estimated coefficients I find on the Chinese leading indicator. I interpret the changes in the means of the in-sample model implied 5y yields, the 5y risk-neutral yields, and the 5y term premia before and after the increase as the average effects of the Chinese leading indicator. These effects are measuring the economic importance of the Chinese leading indicator for the 5y yields, the 5y risk-neutral yields, and the 5y term premia.

My main empirical results yield several new findings.

I find that in the US, a one percentage point lower Chinese leading indicator lowers the 5y Treasury yield and the 5y Treasury term premium by 4.1 basis points over the short run. In the 5th month, the 5y Treasury yield decreases by 10.2 basis points, the 5y Treasury term premium by 9.2 basis points, and the 5y Treasury risk-neutral yield by 1 basis point. The responses of the 5y Treasury yield, the 5y Treasury term premium, and the 5y Treasury risk-neutral yield change by less than 1 basis point (in the absolute terms) in the 12th month. The lower Chinese leading indicator has an economically important negative impact on the 5y Treasury yield and its term premium[7].

At first glance, the Chinese leading indicator affects the 5y Bund yield in the opposite way as the 5y Treasury yield. The lower Chinese leading indicator *increases* the 5y Bund yield and the term premium attached to the 5y German Bunds. Over the short run, the 5y Bund yield increases by 3.8 basis points, the 5y Bund term premium by 3.3 basis points, and the 5y

[7] In my companion paper, Maletic (2018), I find that the lower growth of the Chinese foreign exchange reserves represents incremental information to the Chinese leading indicator and signals a lower 5y Treasury yield and decreases the compensation for bearing the duration risk (the 5y Treasury term premium).

Bund risk-neutral yield by 0.5 basis points. In the 5th month, the 5y Bund yield increases by 5.9 basis points, the 5y Bund term premium by 4.7 basis points, and the 5y Bund risk-neutral yield by 1.2 basis points. However, as such increases are usually associated with recessions the effect could be driven by the ongoing sovereign debt crisis in the euro area. I find that the effects of the Chinese leading indicator on the 5y Bund yield and its term premium change direction and increase in the economic magnitude after the sovereign debt crisis. With the four-factor single country affine term structure model I find that in the 12th month, in the post sovereign debt crisis sample, the model implied 5y Bund yield decreases by 22.5 basis points, the 5y Bund term premium by 21.9 basis points and the 5y Bund risk-neutral yield by 0.6 basis points.

The different direction of the effects in the US and Germany in a joint model after the financial crisis stems from two sources. First, although the principal components are extracted jointly from the US and German yield curves, I condition on a different set of domestic macroeconomic variables when I estimate the average effects of the Chinese leading indicator. Considering the link between the US/German unemployment rates and core inflations, and the difference between the leading indicators of the US/Germany and China is important when quantifying the effect of the Chinese leading indicator on the US and German yield curves. Second, and more importantly, after the financial crisis, we have witnessed the sovereign debt crisis in the euro area. The estimated effects change direction and strengthen in economic magnitude in Germany after the sovereign debt crisis.

My empirical findings suggest that the lower Chinese leading indicator mainly alters the term premia attached to the 5y nominal bonds. It signals lower 5y nominal interest rates and decreases the compensation for bearing the term (duration) risk in the US after the financial crisis, and in Germany after the sovereign debt crisis. The deterioration of the outlook about the Chinese economy provides a signal for lower longer-term nominal interest rates going forward. In an environment with low levels of growth, inflation and accommodative monetary policy constrained with the effective lower bound, investors are willing to accept lower compensation for holding nominal bonds instead of short-term securities, and are very sensitive towards signals about the future growth and inflation risks such as deterioration of the outlook about the Chinese economy.

The rest of this paper is organized as follows. Section 2 introduces an affine term structure model. Section 3 presents the data. Main results are presented in Section 4. Section 5 concludes.

2. Affine Term Structure Model

I estimate an affine term structure model. I use an estimator proposed by Diez de Los Rios (2015, 2018). His asymptotic least-square (ALS) estimator is internally consistent and has a limiting distribution which is asymptotically equivalent to the maximum likelihood. The evolution of the state variables (under the historical measure) follows the vector-autoregressive (VAR) process[8]

$$\begin{bmatrix} X_t^s \\ X_t^u \end{bmatrix} = \mu + \Phi \begin{bmatrix} X_{t-1}^s \\ X_{t-1}^u \end{bmatrix} + \begin{bmatrix} v_t^s \\ v_t^u \end{bmatrix} \tag{1}$$

Where

X_t^s – spanned pricing factors (principal components) $\in \mathcal{R}^{K_s \times 1}$

X_t^u – unspanned macroeconomic variables $\in \mathcal{R}^{K_u \times 1}$

I use the principal component analysis and extract the principal components jointly from the US and German nominal term structures of interest rates (X_t^s). The macroeconomic variables (X_t^u) affect the bond prices merely through the principal components with a lag. In the US, in the model, I include the unspanned macroeconomic variables the US unemployment rate, the US core inflation, the US leading indicator, and the Chinese leading indicator. In Germany, in the model, I include the German unemployment rate, German core inflation, German leading indicator, and the Chinese leading indicator. The principal components which I extract jointly from the US and German yield curves *do not change.*

Shocks, $v_t = [v_t^s \quad v_t^u]'$, conditionally on lagged principal components and unspanned macroeconomic variables follow a Normal distribution, $v_t | \{X_s\}_{s=0}^{t-1} \sim N(0, \Sigma)$. μ, Φ, and Σ are partitioned according to the spanned and unspanned factors. Namely,

---

[8] Adrian, Crump and Moench (2013) were among the first to propose the regression based estimation of an affine term structure model.

$$\mu = \begin{bmatrix}\mu_s\\ \mu_u\end{bmatrix}, \qquad \Phi = \begin{bmatrix}\Phi_{ss} & \Phi_{su}\\ \Phi_{us} & \Phi_{uu}\end{bmatrix}, \qquad and\ \Sigma = \begin{bmatrix}\Sigma_{ss} & \Sigma_{su}\\ \Sigma_{us} & \Sigma_{uu}\end{bmatrix}. \tag{2}$$

The bond pricing factors (principal components) and the nominal short-term interest rates in the US and Germany are related through the affine relation

$$r_{j,t} = \delta_0^{j,s} + \delta_1^{j,s'} X_t^s, \qquad for\ j = US\ and\ Germany. \tag{3}$$

The two-country affine term structure model allows for different loadings ($\delta_0^{j,s}$ and $\delta_1^{j,s'}$) on the US and German nominal short rates. When $\delta_0^{j,s}$ and $\delta_1^{j,s'}$ equal zero for $j = US\ or\ Germany$ the two-country model is reduced to a (usual) single country model[9].

Similarly as in the single-country case, under the risk-neutral probability measure, the spanned and unspanned factors follow the VAR (1) process

$$\begin{bmatrix}X_t^s\\ X_t^u\end{bmatrix} = \begin{bmatrix}\mu_s^\star\\ \mu_u^\star\end{bmatrix} + \begin{bmatrix}\Phi_{ss}^\star & 0\\ \Phi_{us}^\star & \Phi_{uu}^\star\end{bmatrix}\begin{bmatrix}X_{t-1}^s\\ X_{t-1}^u\end{bmatrix} + \begin{bmatrix}v_t^{s\star}\\ v_t^{u\star}\end{bmatrix} \tag{4}$$

Shocks, $v_t{}^\star = [v_t^{s\star} \quad v_t^{u\star}]'$, conditionally on lagged principal components and unspanned macroeconomic variables follow a Normal distribution, $v_t{}^\star|\{X_s\}_{s=0}^{t-1} \sim N(0,\Sigma)$. Σ is the same matrix as in (2). The pricing (risk-neutral) transition matrices, $\mu^\star$ and $\Phi^\star$, can be written as

$$\mu^\star = \begin{bmatrix}\mu_s - \lambda_0^s\\ \mu_u - \lambda_0^u\end{bmatrix} = \begin{bmatrix}\mu_s - \lambda_0^s\\ \mu_u\end{bmatrix} = \begin{bmatrix}\mu_s^\star\\ \mu_u^\star\end{bmatrix}, \qquad \Phi^\star = \begin{bmatrix}\Phi_{ss} - \lambda_1^{ss} & \Phi_{su} - \lambda_1^{su}\\ \Phi_{us} - \lambda_1^{us} & \Phi_{uu} - \lambda_1^{uu}\end{bmatrix} = \begin{bmatrix}\Phi_{ss} - \lambda_1^{ss} & 0\\ \Phi_{us} & \Phi_{uu}\end{bmatrix} = \begin{bmatrix}\Phi_{ss}^\star & 0\\ \Phi_{us}^\star & \Phi_{uu}^\star\end{bmatrix}. \tag{5}$$

Because unspanned macroeconomic variables do not affect bond prices under the pricing measure following Adrian, Crump and Moench (2013), $\lambda_0^u = 0$, $\lambda_1^{us} = 0 \in \mathcal{R}^{K_u\times K_s}$, $\lambda_1^{uu} = 0 \in \mathcal{R}^{K_u\times K_u}$, the upper right $K_s \times K_u$ block of risk-neutral matrix $\Phi^\star$, $\Phi_{su}^\star = (\Phi_{su} - \lambda_1^{su})$ is zero, and therefore $\boldsymbol{\Phi_{su} = \lambda_1^{su}} \in \mathcal{R}^{K_s\times K_u}$.

Given the assumptions (1) – (5), (log) bond prices of maturity n in country j at time period t are exponentially affine in the spanned factors (principal components)

---

[9] $\delta_1^{j,s'}$ is a row vector so it equals a row of zeroes of appropriate dimension.

$$lnP_{j,t}^{(n)} = A_n^{j,s} + B_n^{j,s'} X_t^s \tag{6}$$

The continuously compounded yield on a n-period zero-coupon bond in country j at time t equals $y_{j,t}^{(n)} = -\frac{1}{n} lnP_{j,t}^{(n)}$, and can be written as

$$y_{j,t}^{(n)} = a_n^{j,s} + b_n^{j,s'} X_t^s, \tag{7}$$

where $a_n^{j,s} = -\frac{A_n^{j,s}}{n}$ and $b_n^{j,s} = -\frac{B_n^{j,s}}{n}$.

Following Diez de Los Rios (2018) recursive linear restrictions $A_n^{j,s}$ and $B_n^{j,s'}$ are given as $(for\, n > 1)$

$$A_n^{j,s} = A_{n-1}^{j,s} + B_{n-1}^{j,s}{}'(\mu_s - \lambda_0^s) + \frac{1}{2} B_{n-1}^{j,s}{}' \Sigma_{ss} B_{n-1}^{j,s} - \delta_0^{j,s} \tag{8}$$

$$B_n^{j,s'} = B_{n-1}^{j,s}{}'(\Phi_{ss} - \lambda_1^{ss}) - \delta_1^{j,s'} \tag{9}$$

$$A_0^{j,s} = 0, \quad A_1^{j,s} = -\delta_0^{j,s}, \quad B_0^{j,s'} = 0, \quad B_1^{j,s'} = -\delta_1^{j,s'}, \quad for\, j = US\, and\, Germany. \tag{10}$$

When prices of risk parameters λ_0^s and λ_1^{ss} in (8) and (9) are set to zero, the recursions generate the risk adjusted bond pricing parameters

$$A_n^{j,s,RF} = A_{n-1}^{j,s,RF} + B_{n-1}^{j,s,RF'} \mu_s + \frac{1}{2} B_{n-1}^{j,s,RF'} \Sigma_{ss} B_{n-1}^{j,s,RF} - \delta_0^{j,s} \tag{11}$$

$$B_n^{j,s,RF'} = B_{n-1}^{j,s,RF'} \Phi_{ss} - \delta_1^{j,s'} \tag{12}$$

Risk-adjusted parameters imply that the model-fitted yields equal the time t expectation of the average future short rates over the next n periods, $E_t\left(-\left(\frac{1}{n}\right) lnP_{j,t}^{(n)}\right) = -\left(\frac{1}{n}\right)\left(A_n^{j,s,RF} + B_n^{j,s,RF'} X_t^s\right)$. The risk neutral yield (RNY), and the term premium (TP), the difference between the model-implied fitted yield and the risk neutral yield, can be written as[10,11]

[10] Campbell, Sunderam and Viceira (2009), Christensen, Lopez and Rudebusch (2010), Hördahl and Tristani (2012), and Rousselett (2017), amongst others, investigate the importance of variation in the estimated term premium for long-term nominal interest rates. They decompose the model implied term premium of the long-term nominal interest rates into the real term premium and the inflation risk premium. Abrahams, Adrian, Crump,

$$RNY_{j,t}^{(n)} = -\left(\frac{1}{n}\right)\left[A_n^{j,s,RF} + B_n^{j,s,RF'} X_t^s\right] \tag{13}$$

$$TP_{j,t}^{(n)} = -\left(\frac{1}{n}\right)\left[\left(A_n^{j,s} - A_n^{j,s,RF}\right) + \left(B_n^{j,s} - B_n^{j,s,RF}\right)' X_t^s\right] \tag{14}$$

Diez de Los Rios (2018) notices that when the state variables are linear combinations of yields (i.e., $X_t^s = P' y_{j,t}^{(n)} = P'\left(a_n^{j,s} + b_n^{j,s'} X_t^s\right)$, for some full-rank matrix P) self-consistency implies[12,13]

$$P'a(\theta) = 0, \qquad P'b(\theta) = I,$$

where $\theta = (\theta_1', \theta_2', \theta_3')'$, $\theta_1 = vec(\Theta^\star)$, $\theta_2 = vec[(\mu\ \Phi)']$, $\theta_3 = vech(\Sigma^{1/2})$, and

$$\Theta^{\star\prime} = \begin{pmatrix} \delta_0^{US,s} & \delta_1^{US,s'} \\ \delta_0^{GER,s} & \delta_1^{GER,s'} \\ \mu_s^\star & \Phi_{ss}^\star \end{pmatrix}. \tag{15}$$

Diez de Los Rios (2015) exploits conditions in (15) and proposes an asymptotic least squares (ALS) estimator. Estimator in Diez de Los Rios (2018) allows estimation of a multi-country affine term structure model with a large number of spanned factors (principal components).

To investigate how the Chinese leading indicator affects the spanned factors (X_t^s) and (log) bond prices ($lnP_{j,t}^{(n)}$) I focus on $\hat{\lambda}_{su}$. I increase principal components extracted from the US and German yield curves by the estimated coefficients $\hat{\lambda}_{su}$ which are statistically significantly different from zero and correspond to the Chinese leading indicator. I compare the change in the mean of the model implied 5y Treasury/Bund yields, the 5y Treasury/Bund risk-neutral

---

Moench and Yu (2016) show that announcements of asset purchase programmes lower the long-term nominal interest rates mainly by lowering the model implied real term premium.

[11] Bernanke (2015) points out that after 2013 the 10-year Treasury term premium is more important for low 10-year Treasury yield than the 10-year Treasury risk-neutral yield.

[12] Cochrane and Piazzesi (2005) pointed out that variables which are linear combinations of yields, state variables which come out of the model, should be equal to imposed observed pricing factors.

[13] To ensure the positivity of covariance matrix Σ Diez de Los Rios (2018) focuses on its Cholesky decomposition, $\Sigma = \Sigma^{1/2}\Sigma^{1/2'}$.

yields, and the 5y Treasury/Bund term premia before and after I increase the principal components by the estimated coefficients $\hat{\lambda}_{su}$. I interpret the difference in the means as the average effect of the Chinese leading indicator on the model implied 5y Treasury/Bund yields, the model implied 5y Treasury/Bund risk-neutral yields, and the model implied 5y Treasury/Bund term premia.

3. Data

I estimate the joint model of the US and German nominal term structure of interest rates in the post financial crisis sample, from June 2009 to December 2017. The parameters of the zero-coupon yield curve are retrieved from Deutsche Bundesbank (BUBA) and Gürkaynak, Sack and Wright (2007).

I focus on the maturities from 1 to 60 months (5 years). The rest of the data is as follows. Core inflation and unemployment rates for the US and Germany are from the FRED database of the Federal Reserve Bank of St. Louis and from Eurostat. I retrieve the leading indicators of the US, German and Chinese economies from OECD[14].

The leading indicator is constructed in such a way as to identify and predict the turning points in the business cycles. Reference series which is chosen to approximate the economic activity is the quarterly growth of the GDP. The OECD generates monthly estimates of the GDP based on the official quarterly estimates.

The database on the main economic indicators (MEI) provides the main source of variables that are included in the indicator. The variables can be grouped in (1) GDP and industrial production, (2) selected commodity output variables (crude steel, crude petroleum etc.), (3) business and consumer tendency survey series, (4) selected manufacturing variables (deliveries, stocks, new orders etc.), (5) construction, (6) domestic trade, (7) labor market series, (8) consumer and producer prices, (9) money aggregates, (10) interest rates, (11) financial variables, (12) exchange rates, (13) international trade and (14) balance of payments data.

[14] Available at http://www.oecd.org/std/leading-indicators/

I use the trend restored version of the index in 12-month log differences. This version of the index most closely tracks the yearly GDP growth rate and is available at a monthly frequency.

Figure 2 depicts the six principal components extracted jointly from the US and German term structures in the post financial crisis sample. The loadings on the yields of different maturities of the first and the second principal component are a mixture of level and slope. Loadings of the third to the sixth principal component do not have meaningful economic interpretation. That is why I decided to focus instead on the time-series dimension.

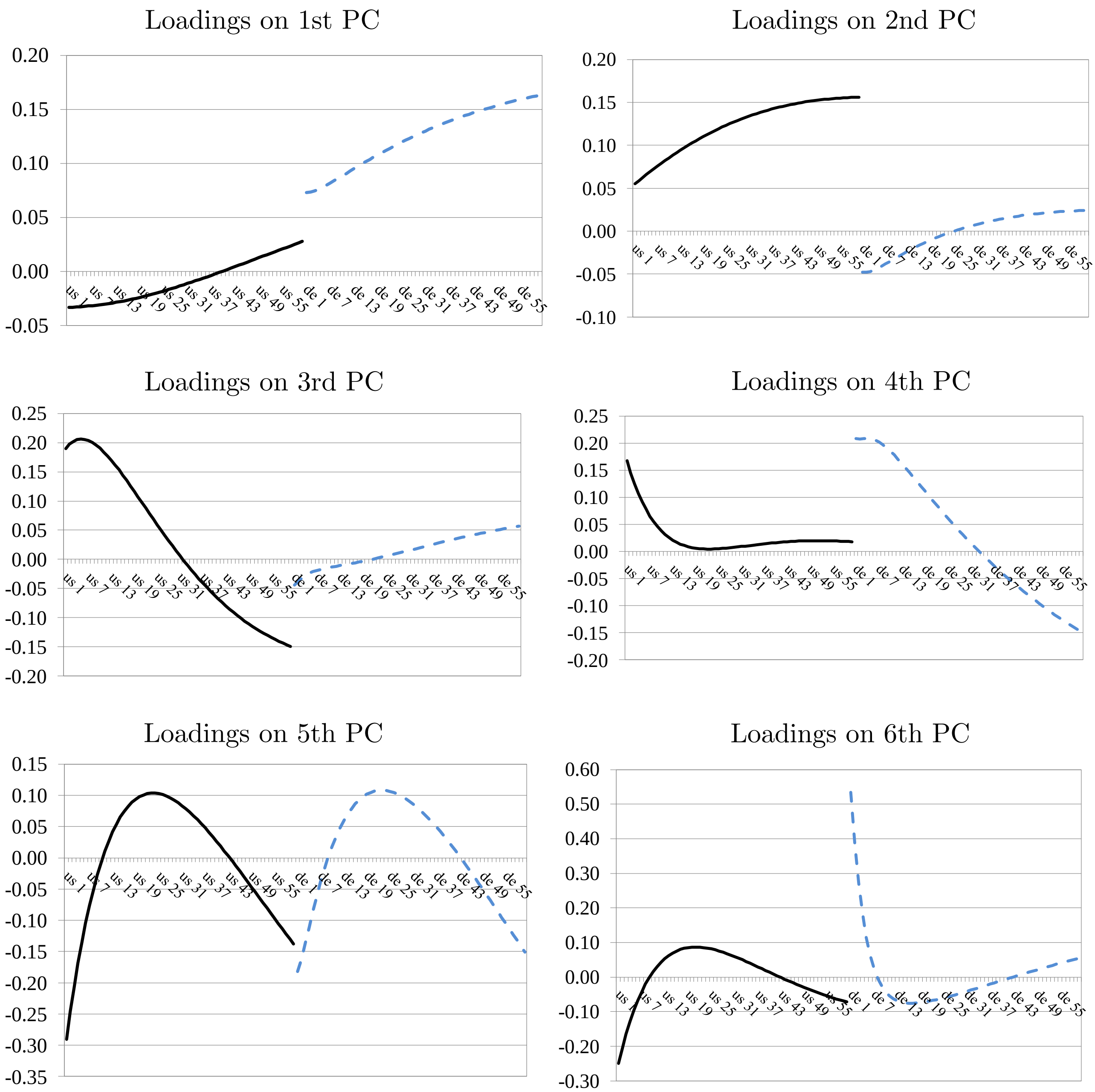


Figure 2: Loadings of US (black line) and German (dashed-blue line) monthly zero-coupon yields with maturities of one to sixty months (5 years) on the six global principal components. Sample spans from June 2009 to December 2017.

Figure 3 plots the first two principal components extracted jointly from the US and German nominal term structures in the post financial crisis sample. In the post financial crisis sample, the dynamics of principal component 1 are similar to the dynamics of the 5y Bund yield. The dynamics of principal component 2 are similar to the dynamics of the 5y Treasury yield. While principal components 1 and 2 do not have meaningful interpretations in the cross-sectional dimension (both are a mixture of levels and slopes), the principal components 1 and 2 in the time-series dimension show clear diverging patterns which are most probably due to the different monetary policy stances in the US and Germany in the post financial crisis sample.

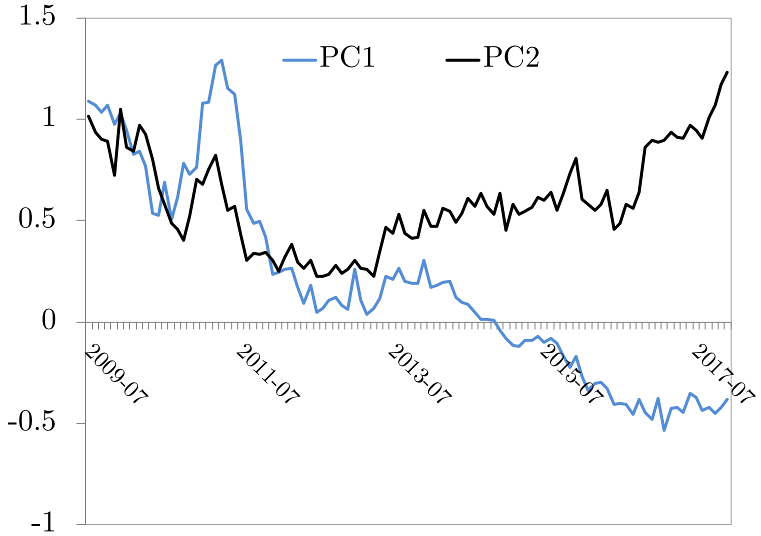


Figure 3: Principal component 1 and 2 in the post financial crisis sample (from June 2009 to December 2017).

Table 1 shows the average percentage of explained variation of 60 yields with monthly maturities when I use one to six principal components. One factor model shows a clear disconnect between the US and German yields. While the first principal component extracted jointly from the US and German nominal term structures explains 97 percent of Bund yield variation, it explains only 11 percent of Treasury yield variation (up to the maturity of 5 years). The two-factor model already explains almost 90 percent of the variation of the US and more than 98 percent of the variation of German yields. However, the pattern of loadings on the yields of different maturities of the first two principal components in the US and Germany is not clear. Figure 4 depicts R^2s of the first six principal components on 120 yields with monthly maturities, 60 in the US and 60 in Germany. Model almost fully explains the yield variation.

Table 1: Average percentage of explained variation of 60 monthly maturity yields in the US and Germany when I use one, two, three, four, 5, or six principal components from June 2009 to December 2017.

	One Factor	Two Factors	Three Factors	Four Factors	Five Factors	Six Factors
U.S.	11.2%	89.8%	98.8%	99.2%	99.8%	99.9%
Germany	97.0%	98.4%	98.5%	99.6%	99.8%	99.9%

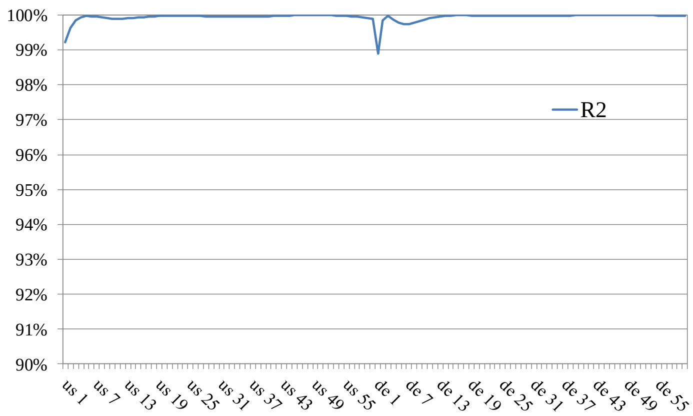


Figure 4: Percentage of explained variation of monthly yields with maturities of one to sixty months (5 years) in the US and Germany with the global six-factor model (which uses global PC1 to PC6). Sample spans from June 2009 to December 2017.

Figure 5 depicts the growth of the Chinese leading indicator before and after the financial crisis. The average growth of the Chinese leading indicator from 1998 to 2007, 14.2 percent, decreased to 10.3 percent in the post financial crisis sample. The growth of Chinese leading indicator after the financial crisis exhibits a clear downward trend. The yearly growth of the leading indicator in December 2017 decreased to 5 percent. From June 2009 to December 2017 the mean of the Chinese leading indicator is equal to 10.3 percent, and its standard deviation is equal to 4.5 percent.

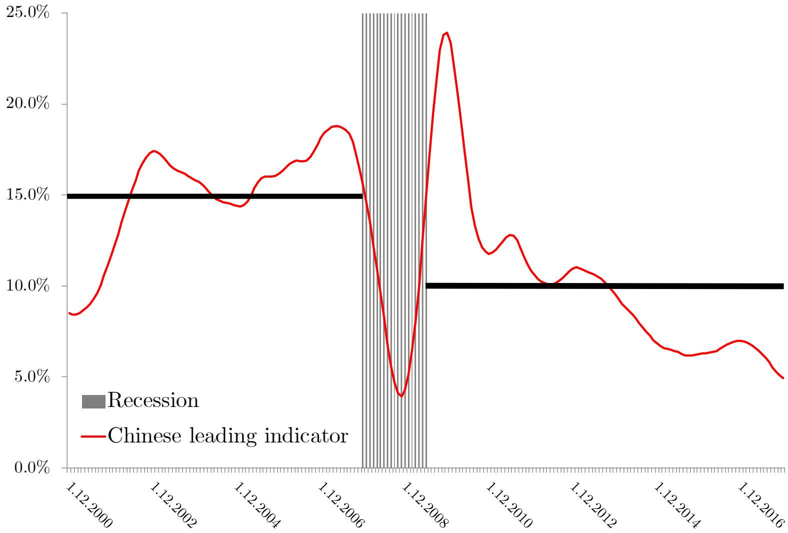


Figure 5: Average growth of the Chinese composite leading indicator before and after the financial crisis. Twelve-month log differences. Sample spans from December 2000 to December 2017. Source: OECD.

Figure 6 shows the US, Chinese and German leading indicators in the post financial crisis sample. Average yearly growth rates of the US and German leading indicators are similar. From June 2009 to December 2017, on average, US leading indicator increased by 2.1 percent. The German leading indicator increased by 2 percent. However, the German leading indicator seems to be exhibiting larger cyclical movements than the US leading indicator in the post financial crisis sample. Its standard deviation is 2.2 percent compared to 1.4 percent in the US. The growth of the Chinese leading indicator is converging towards the growth of German and the US leading indicator.

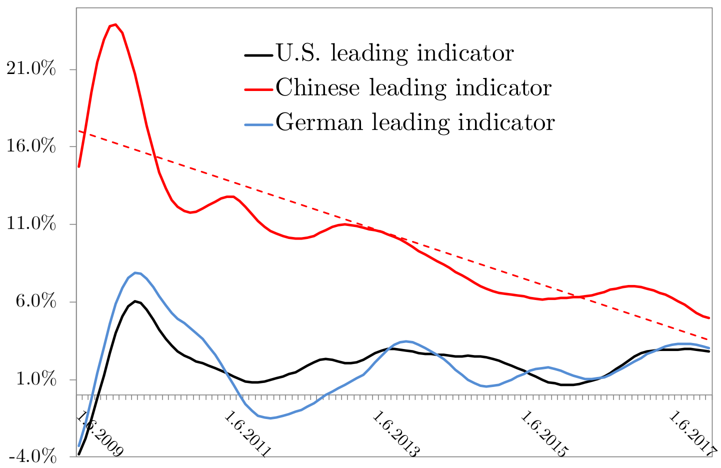


Figure 6: The US, German and Chinese composite leading indicators (CLIs) with trends-restored. Twelve-month log differences. Sample spans from June 2009 to December 2017. Source: OECD.

Figure 7 (left panel) presents the US and German core inflations. In the post financial crisis sample, the average German core inflation equals 1.1 percent. The average core inflation in the US equals 1.7 percent. Core inflations are below 2 percent, the policy target inflation rate. Figure 7 (right panel) shows the unemployment rates. In the US the unemployment rate decreased from 10 percent in September 2009 to 4.1 percent by December 2017. The German unemployment rate decreased from 7.9 percent in July 2009 to 3.6 percent by December 2017.

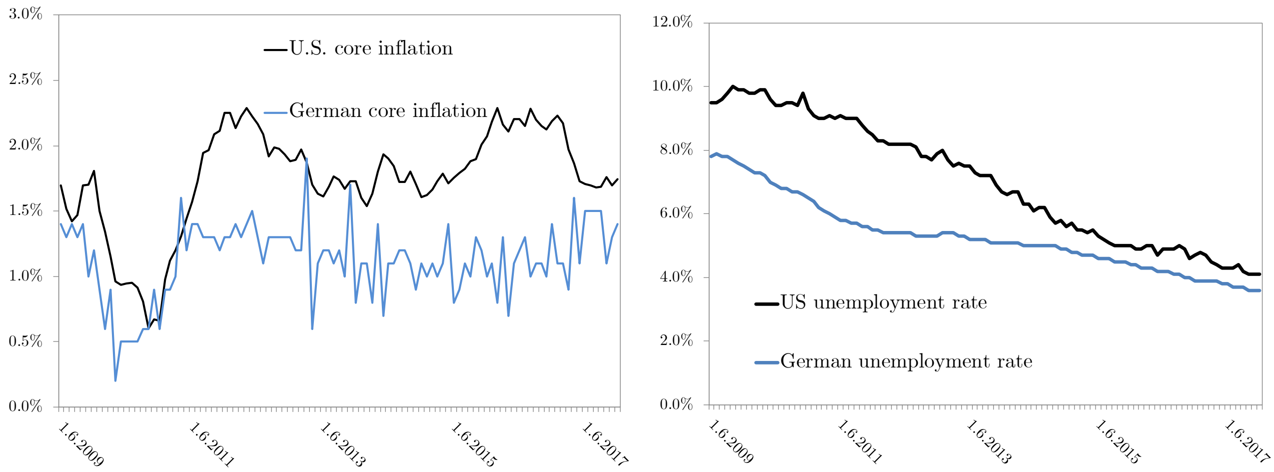


Figure 7: US and German core inflation rates (left panel) and unemployment rates (right panel). Sample spans from June 2009 to December 2017. Source: St. Louis FRED and Eurostat.

In Figure 8 we can see that the 5y Bund term premium decreased from 2.4 percent in December 2009 to 1.2 percent by August 2010. The Chinese leading indicator decreased from 24 percent to 13 percent over the same period. During the sovereign debt crisis, the 5y Bund term premium temporarily increased to 2.5 percent in March 2011 but decreased to 80 basis points by December 2011. By September 2016, the 5y Bund term premium decreased to −40 basis points. It increased to 28 basis points by December 2017. The Chinese leading indicator, on the other hand, steadily decreased from 13 percent in August 2010 to 5 percent by December 2017. The in-sample correlation of the two series equals 0.86.

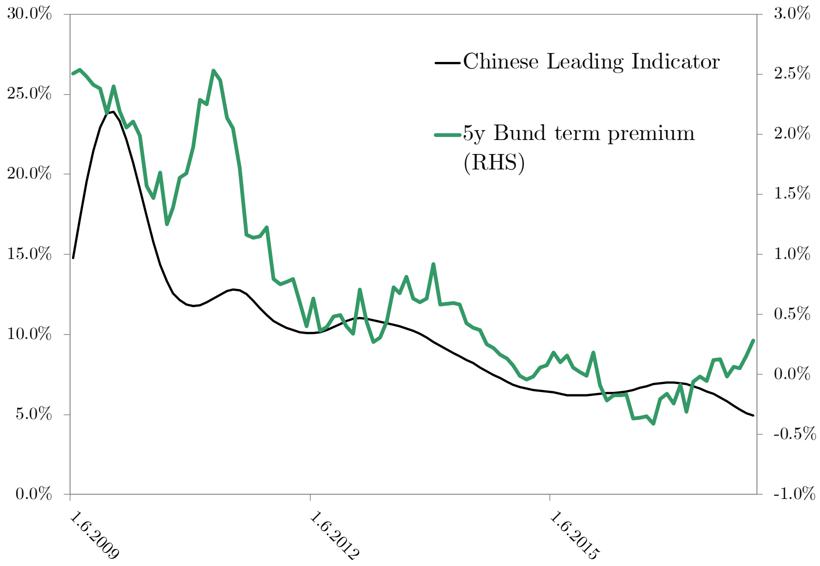


Figure 8: 5y Bund term premium and the Chinese leading indicator. Sample spans from June 2009 to December 2017. Source: OECD.

Table 2 presents descriptive statistics. After the crisis, the average German unemployment rate, 5.3 percent, is smaller than in the US, 7 percent. Average German nominal short rate is negative, –4 basis points. Its standard deviation is higher than in the US, 49 compared to 34 basis points. In the post financial crisis sample, the average 5-year Bund yield equals 69 basis points and is lower than in the US, 154 basis points.

Table 2: Descriptive statistics. Yearly growth of OECD leading indicators, core inflation rates, unemployment rates, nominal short rates (one-month government nominal yields), and the 5-year government nominal yields in the US and Germany. Sample spans from June 2009 to December 2017. Source: FRED database of the Federal Reserve Bank of St. Louis, ECB, Eurostat, OECD, Deutsche Bundesbank (BUBA).

	Mean		Standard deviation		Percentiles			
					US		GER	
	US	GER	US	GER	5th	95th	5th	95th
OECD leading indicator	2.1%	2.0%	1.4%	2.2%	0.7%	4.8%	-1.3%	6.9%
Core Inflation	1.7%	1.1%	0.4%	0.3%	0.9%	2.2%	0.6%	1.5%
Unemployment rate	7.0%	5.3%	1.9%	1.1%	4.3%	9.8%	3.7%	7.6%
Short rate	39	-4	34	49	3	125	-82	79
5-year yield	154	69	51	98	70	240	-50	250

4. Main Results

In this section, I present my main empirical results. Before I introduce the decomposition of the 5y Treasury and Bund yields in the 5y Treasury/Bund risk-neutral yields and the 5y Treasury/Bund term premia, I present the short and long-run effects of the Chinese leading indicator on the actual 5y Treasury/Bund yields.

Table 3 presents the estimated coefficients of the five-variable vector-autoregression for the US economy. All variables included in the regression are in percentage points. Estimated effects are for a percentage point increase, except for the Chinese leading indicator where the estimated effects are for a percentage point decrease. In the last column of Table 3, we can see that a percentage point lower Chinese leading indicator on average decreases the 5y Treasury yield by 4.3 basis points. In the first row of Table 3, we can see that the effects of the US macroeconomic variables are significant. A percentage point increase of the US unemployment on average decreases the 5y Treasury yield by 12.2 basis points. Higher US core inflation on average decreases the 5y Treasury yield by 28.9 basis points. Although significant at the 10% level, the effect of the US leading indicator is small in economic magnitude, on average 2.7 basis points.

Table 3: Estimated coefficients of a five-variable vector autoregression: $X_t = \mu + \Phi X_{t-1} + \varepsilon_t$. Variables included in the regression: 5y Treasury yield, US unemployment, US core inflation, US leading indicator, and Chinese leading indicator. Sample spans from June 2009 to December 2017. Bolded coefficients are significant at the 10% level.

Factor	$\boldsymbol{\Phi_{1,1}}$	$\boldsymbol{\Phi_{1,2}}$	$\boldsymbol{\Phi_{1,3}}$	$\boldsymbol{\Phi_{1,4}}$	$\boldsymbol{\Phi_{1,5}}$
	(5y Treasury Yield)	(ur_{us})	($CCPI_{us}$)	(CLI_{us})	(CLI_{ch})
5y Treasury Yield	**0.7092**	**-0.1222**	**-0.2890**	**-0.0269**	**0.0425**
(t-statistic)	11.21	-3.87	-3.50	-1.72	3.36
ur_{us}	-0.0607	**0.9438**	-0.0950	**-0.0311**	**0.0244**
(t-statistic)	-1.28	40.04	-1.54	-2.66	2.58
$CCPI_{us}$	0.0025	0.0103	**0.9131**	**-0.0224**	-0.0088
(t-statistic)	0.08	0.62	20.93	-2.71	-1.32
CLI_{us}	-0.0630	**-0.2065**	0.1231	**0.8412**	**0.1326**
(t-statistic)	-1.10	-7.20	1.64	59.18	11.56
CLI_{ch}	-0.0173	**-0.1374**	**-0.3702**	**-0.4144**	**1.0544**
(t-statistic)	-0.19	-3.04	-3.13	-18.48	58.28

Table 4 present the results for the German economy. Macroeconomic variables affect the 5y Bund yield in the opposite direction than in the US. A percentage point higher German unemployment increases the 5y Bund yield, on average by 19 basis points. The effect of German core inflation is positive but becomes insignificant. This suggests that in Germany, the 5y Bund yield was reacting more to the output gap than to inflation in the post financial crisis sample.

Estimated coefficients on the leading indicators are significant but of the opposite sign than in the US. A percentage point higher German leading indicator increases 5y Bund yield on average by 2.3 basis points. The economic magnitude is fairly similar to the Chinese leading indicator. A percentage point lower Chinese leading indicator increases the 5y Bund yield on average by 2.5 basis points.

Table 4: Estimated coefficients of a five-variable vector autoregression: $X_t = \mu + \Phi X_{t-1} + \varepsilon_t$. Variables included in the regression: 5y Bund yield, German unemployment, German core inflation, German leading indicator, and Chinese leading indicator. Sample spans from June 2009 to December 2017. Bolded coefficients are significant at the 10% level.

Factor	$\boldsymbol{\Phi_{1,1}}$	$\boldsymbol{\Phi_{1,2}}$	$\boldsymbol{\Phi_{1,3}}$	$\boldsymbol{\Phi_{1,4}}$	$\boldsymbol{\Phi_{1,5}}$
	(5y Bund Yield)	(ur_{ger})	($CCPI_{ger}$)	(CLI_{ger})	(CLI_{ch})
5y Bund Yield	**0.8568**	**0.1915**	0.1145	**0.0226**	**-0.0247**
(t-statistic)	18.80	3.22	1.48	2.17	-2.27
ur_{ger}	-0.0163	**0.9954**	-0.0064	**-0.0079**	0.0018
(t-statistic)	-1.13	52.68	-0.26	-2.40	0.53
$CCPI_{ger}$	**0.2839**	**-0.4714**	**-0.1853**	**-0.0988**	**0.0539**
(t-statistic)	5.07	-6.46	-1.95	-7.74	4.04
CLI_{ger}	**-0.5205**	0.1902	**0.3537**	**0.9278**	**0.1122**
(t-statistic)	-5.76	1.61	2.30	44.98	5.20
CLI_{ch}	0.0595	0.3001	**0.4879**	**-0.1329**	**0.9443**
(t-statistic)	0.50	1.93	2.41	-4.88	33.19

Figure 9 depicts the impulse response functions of the 5y Treasury yield (left panel) to one percentage point negative shock to the Chinese leading indicator. The response of the 5y Treasury yield strengthens from -4.2 basis points in the 1^{st} month to -10.6 basis points in the 6^{th} month. Afterwards, it reverts to -7.6 basis points and remains significantly different from 0 in the 12^{th} month. In the right panel of Figure 9, we can observe that the economic magnitude of the response of the 5y Bund yield is smaller and goes in the opposite way than in the US. The response is significant only in the first month. The 5y Bund yield increases by 2.5 basis points.

Response of the 5y Treasury yield to 1% shock to Chinese leading indicator (orthagonalized IRF) (in bps)

Response of the 5y Bund yield to 1% shock to Chinese OECD leading indicator (orthagonalized IRF) (in bps)

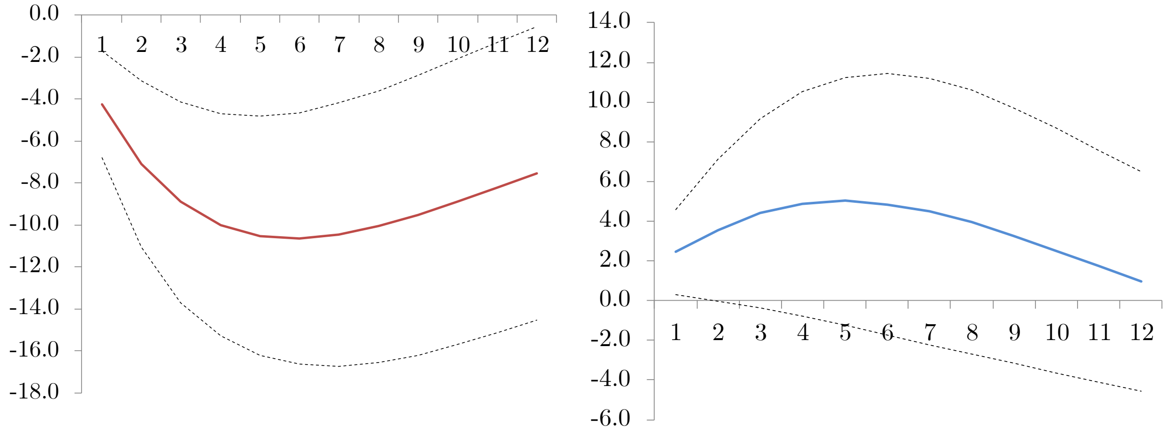


Figure 9: Orthogonalized impulse response functions of the 5y Treasury yield (left panel) and the 5y Bund yield (right panel) to 1 percentage point shock to Chinese OECD leading indicator. I estimate two structural vector auto-regressions with Cholesky identification scheme. Sample spans from June 2009 to December 2017. Variables included in the US model: 5y Treasury yield, US unemployment, US core inflation, US leading indicator, and Chinese leading indicator. Variables included in the German model: 5y Bund yield, German unemployment, German core inflation, German leading indicator, and Chinese leading indicator. Chinese leading indicator is ordered last and lower-triangular variance-covariance matrix of shocks is imposed.

Next, I estimate the two-country affine term structure model with the unspanned macroeconomic variables following Diez de Los Rios (2018). I use principal component analysis and extract principal components which are explaining the most of variation in the US and German yield curves. In the US, in the model, I include the six principal components, the US unemployment, US core inflation rate, the US leading indicator, and the Chinese leading indicator. In Germany, in the model, I include German unemployment, German core inflation, German leading indicator, and the Chinese leading indicator. I estimate the affine term structure model with the unspanned macroeconomic variables. I use the identity $\hat{\Phi}_{su} = \hat{\lambda}_1^{su}$ to measure the effects of the macroeconomic variables on the six principal components, and bond prices.

Table 5 presents the estimated prices of risks of the six-factor model for the US economy. Average pricing error shrinks from 4.3 basis points to 1.4 basis points as I move from the five to the six-factor model (of the fitted 5y Treasury yield). The risk of the first principal component is affected significantly by itself, the sixth principal component, the US unemployment, US core inflation, and the US leading indicator. Risks of the second principal component are affected significantly by the second, the third principal component and all four macroeconomic variables: the US unemployment, US core inflation, US leading indicator, and the Chinese leading indicator. The risks of the third principal component are affected significantly by the first principal component, the second, the third and the fourth principal components. The US unemployment, US core inflation, and the Chinese leading indicator affect the risk of the third principal component significantly. The fourth principal component is affected significantly by the first, the fourth principal component, the US leading indicator, and the Chinese leading indicator. The fifth PC is affected significantly by itself and the US leading indicator. The sixth PC is affected significantly by the third, the fourth, the sixth PC, and the US and Chinese leading indicators.

Table 5: Estimated prices of risk, λ_0^s and λ_1^s, of the two-country affine term structure model using an estimator as outlined in Diez de Los Rios (2018). Sample spans from June 2009 to December 2017. Spanned factors: $X_t^s = [PC\ 1_t \quad PC\ 2_t \quad PC\ 3_t \quad PC\ 4_t \quad PC\ 5_t \quad PC\ 6_t]'$. Unspanned factors: $X_t^u = [ur_{US,t} \quad CCPI_{US,t} \quad CLI_{US,t} \quad CLI_{CH,t}]'$. Bolded coefficients are significant at the 10% level. I present the remaining estimated parameters in Appendix A.1. $ur_{US,t}$ – US unemployment rate, $CCPI_{US,t}$ – US core inflation rate, $CLI_{US,t}$ – US leading indicator, $CLI_{CH,t}$ – Chinese leading indicator.

Factor	$\boldsymbol{\lambda_0}$	$\boldsymbol{\lambda_{1,1}}$	$\boldsymbol{\lambda_{1,2}}$	$\boldsymbol{\lambda_{1,3}}$	$\boldsymbol{\lambda_{1,4}}$	$\boldsymbol{\lambda_{1,5}}$	$\boldsymbol{\lambda_{1,6}}$	$\boldsymbol{\lambda_{1,7}}$	$\boldsymbol{\lambda_{1,8}}$	$\boldsymbol{\lambda_{1,9}}$	$\boldsymbol{\lambda_{1,10}}$
	(constant)	(PC1)	(PC2)	(PC3)	(PC4)	(PC5)	(PC6)	(ur_{us})	($CCPI_{us}$)	(CLI_{us})	(CLI_{ch})
PC 1	-0.0534	**-0.3268**	0.0464	-0.2144	0.0532	0.0530	**-0.9977**	**5.8453**	**-16.4305**	**-2.5910**	-0.1548
(t-statistic)	-0.251	-3.721	0.666	-1.605	0.161	0.128	-1.721	2.172	-4.507	-2.789	-0.236
PC 2	**0.5769**	-0.0391	**-0.2498**	**0.3264**	-0.1057	-0.4421	-0.8609	**-5.7030**	**-11.7262**	**-1.5982**	**1.7375**
(t-statistic)	2.656	-0.436	-3.511	2.389	-0.313	-1.046	-1.452	-2.077	-3.153	-1.686	2.594
PC 3	**-0.2106**	**-0.0493**	**0.0768**	**-0.0869**	**0.2320**	-0.0539	-0.0102	**2.5437**	**1.8737**	0.2622	**-0.5069**
(t-statistic)	-3.461	-1.958	3.757	-2.155	2.372	-0.434	-0.057	3.348	1.821	1.000	-2.735
PC 4	0.0973	**0.0524**	-0.0213	0.0203	**-0.4760**	-0.1031	0.1932	-0.6581	-0.7052	**0.8485**	**-0.6093**
(t-statistic)	1.649	2.144	-1.071	0.517	-4.998	-0.851	1.107	-0.895	-0.708	3.341	-3.394
PC 5	-0.0071	-0.0087	-0.0014	-0.0011	0.0528	**-0.2402**	-0.1820	-0.0269	-0.4523	**-0.5021**	0.0593
(t-statistic)	-0.195	-0.568	-0.108	-0.042	0.843	-2.941	-1.521	-0.060	-0.749	-3.264	0.546
PC 6	0.0163	0.0110	-0.0060	**0.0445**	**-0.1970**	0.1134	**-0.1896**	0.0338	0.1168	**0.3071**	**-0.1999**
(t-statistic)	0.454	0.738	-0.506	1.925	-3.471	1.586	-1.862	0.075	0.191	1.969	-1.813

Figure 10 depicts the estimated 5y Treasury risk-neutral yield (left panel) and 5y Treasury term premium (right panel). The 5y Treasury risk-neutral yield increased from 47 basis points in June 2009 to 3.4 percent by December 2017. From December 2014 to December 2017 the 5y Treasury risk-neutral yield increased from -5 basis points to 3.4 percent. The 5y Treasury term premium (upper right panel, Figure 10) decreased from 2.1 percent in June 2009 to 1 percent by October 2010, from where it increased to 2.2 percent in March 2011. Volatile 5y Treasury term premium can be at least in some part explained by the ongoing sovereign debt crisis in the Euro Area. The 5y Treasury term premium decreased to 71 basis points by April 2013 from where it increased to 2 percent in December 2013. After 2014, the 5y Treasury term premium decreased to -1.2 percent by December 2017.

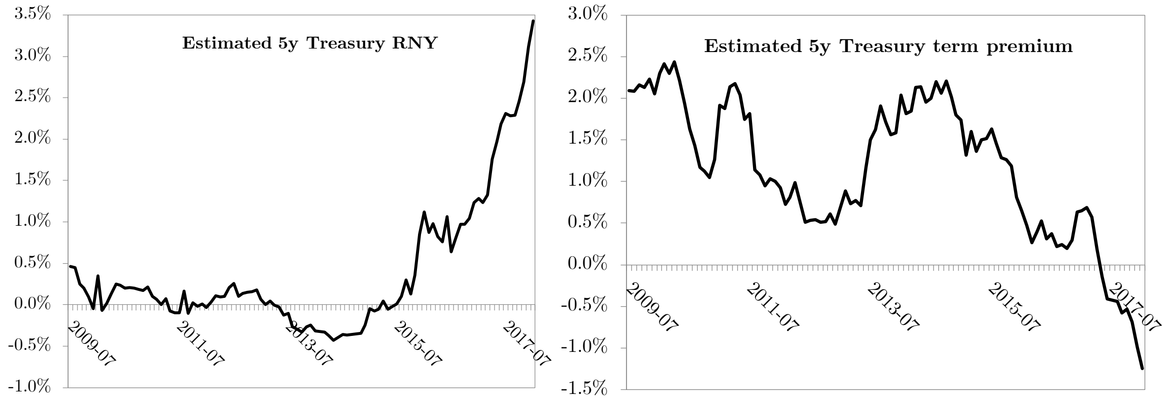


Figure 10: Model implied 5-year Treasury risk-neutral yield (expected future nominal short rate) (left panel) and the 5y Treasury term premium (right panel) estimated with the six-factor model (which uses PC1 to PC6) and unspanned macroeconomic variables: US unemployment, US core inflation, US leading indicator, and the Chinese leading indicator. I use an estimator as outlined in Diez de Los Rios (2018). Sample spans from June 2009 to December 2017.

To measure the effect of the Chinese leading indicator on the US yield curve, I increase the principal components by the significant coefficients $\hat{\lambda}_{1,10}$ which are estimated in Table 5. I calculate the change of the in-sample mean of the 5y Treasury yield, the 5y Treasury risk-neutral yield and the 5y Treasury term premium which are presented in Figure 10. The mean of the model implied 5y Treasury yield increases by 4.1 percent. The mean of the 5y Treasury term premium increases by 4.1 percent and the mean of the model implied 5y Treasury risk-neutral yield remains unchanged. However, the estimated coefficients represent an increase of the Chinese leading indicator by “a unit”. This implies that the Chinese leading indicator would increase by 100 percent. I divide the estimated effects by 100 and premultiply them by -1.

A one percentage point decrease of the Chinese leading indicator decreases the model implied 5y Treasury yield by 4.1 basis points and the model implied 5y Treasury term premium by 4.1 basis points while leaving the model implied 5y Treasury risk-neutral yield unchanged[15].

The 95 percent confidence interval (in basis points) for a percentage points decrease of the Chinese leading indicator of the model implied 5y Treasury yield is $(-7.1,\ -1)$, the model implied 5y Treasury term premium is $(-6.9,\ -1.3)$ and of the model implied 5y Treasury risk-neutral yield is $(-0.2,\ 0.3)$.

Figure 11 presents the impulse response functions of the six principal components to "a unit" shock to the Chinese leading indicator in the model which includes the US macroeconomic variables. The estimated effect on the second principal component, 1.7375 (estimated coefficient in the 10^{th} column of Table 5) decreases to 0.0027 (upper-middle panel in Figure 11, the effect is pre-multiplied by -1). The response of the second principal component increases to 0.0065 by the 5^{th} month from where it reverts back to 0. In Figure 11, we can observe that the responses of the third, the fourth, and the fifth principal components are significant as well.

To measure the long-run effects of the Chinese leading indicator on the US yield curve, I proceed as follows. First, I compute average 5y Treasury yield, average 5y Treasury term premium and average 5y Treasury risk-neutral yield with the principal components ($\mathbf{X}^{\mathbf{s}}$) which I extract from the yield curves. Second, I increase the principal components by responses in the 5th month when the responses of the 2nd, the 3rd, and the 4th PCs are significant and the highest (please refer to Figure 11). Third, I re-compute average 5y

---

[15] When I do not condition on the US unemployment, US core inflation and the US leading indicator, average estimated effects in the US decrease to (in absolute terms): the model implied 5y Treasury yield decreases by 0.8 basis points, the model implied 5y Treasury term premium by 0.6 basis points, and the model implied 5y Treasury risk-neutral by 0.2 basis points.

Treasury yield, average 5y Treasury term premium and average 5y Treasury risk-neutral yield with the new principal components.

In particular, let P denote the Cholesky decomposition of Σ, such that $\mathrm{PP}' = \Sigma$. Furthermore, let $\mathrm{u_t} \in \mathcal{R}^{\mathrm{K}\times 1}$ be such that I can write shocks in (1) as $\mathrm{u_t} = \mathrm{P}^{-1}\mathrm{v_t}$. Orthogonalized impulse responses can be written as[16]

$$\begin{aligned}
\begin{bmatrix} \widehat{\mathrm{X}}_0^{\mathrm{s}} \\ \widehat{\mathrm{X}}_0^{\mathrm{u}} \end{bmatrix} &= \mathrm{I_K}\widehat{\mathrm{P}} \\
\begin{bmatrix} \widehat{\mathrm{X}}_1^{\mathrm{s}} \\ \widehat{\mathrm{X}}_1^{\mathrm{u}} \end{bmatrix} &= \widehat{\Phi}\widehat{\mathrm{P}} \\
\begin{bmatrix} \widehat{\mathrm{X}}_2^{\mathrm{s}} \\ \widehat{\mathrm{X}}_2^{\mathrm{u}} \end{bmatrix} &= \widehat{\Phi}^2\widehat{\mathrm{P}} \\
&\vdots \\
\begin{bmatrix} \widehat{\mathrm{X}}_{\mathrm{T}}^{\mathrm{s}} \\ \widehat{\mathrm{X}}_{\mathrm{T}}^{\mathrm{u}} \end{bmatrix} &= \widehat{\Phi}^{\mathrm{T}}\widehat{\mathrm{P}}
\end{aligned} \tag{16}$$

To scale the orthogonalized impulse of the Chinese leading indicator to a unit shock, I divide the responses by the standard deviation of the Chinese leading indicator (which I order last). The standard deviation corresponds to the element in the last row of the last column of $\widehat{\mathrm{P}}$. I multiply the responses by 10.000 to scale them to basis points responses.

Next, I collect significant responses of the 6 principal components in the fifth month in a row vector which I denote ϕ_{LR}, and add the ϕ_{LR} to the principal components, which I extract from the yield curves

$$\widetilde{\mathrm{X}}_{\mathrm{t}}^{\mathrm{s}} = \mathrm{X}_{\mathrm{t}}^{\mathrm{s}} + \widehat{\phi}_{\mathrm{LR}} \tag{17}$$

I re-estimate the 5y yield, the term premium and the risk-neutral yield with the new principal components, $\tilde{X}_t^s$. I interpret the changes in the 5y yield, the term premium and the risk-neutral yield, which are estimated with X_t^s and $\tilde{X}_t^s$, as average effects of a one percentage

[16] And interpreted as one standard-deviation impulse to u_t.

point increase of the Chinese leading indicator on the 5y yield, the term premium, and the risk-neutral yield.

In the 5^th month, a one percentage point decrease of the Chinese leading indicator decreases the in-sample average of the model implied 5y Treasury yield by 10.2 basis points, the 5y Treasury term premium by 9.2 basis points and the 5y Treasury risk-neutral yield by 1 basis point. In the 12^th month, the model implied 5y Treasury yield decreases by 0.77 basis points, the 5y Treasury term premium by 0.33 basis points and the 5y Treasury risk-neutral yield by 0.44 basis points.

The 95 percent confidence interval (in basis points) for a percentage points decrease of the Chinese leading indicator of the model implied 5y Treasury yield in the 5^{th} month is $(-17, -3.5)$, the model implied 5y Treasury term premium is $(-16.6, -1.9)$ and of the model implied 5y Treasury risk-neutral yield is $(-0.4, -1.6)$.

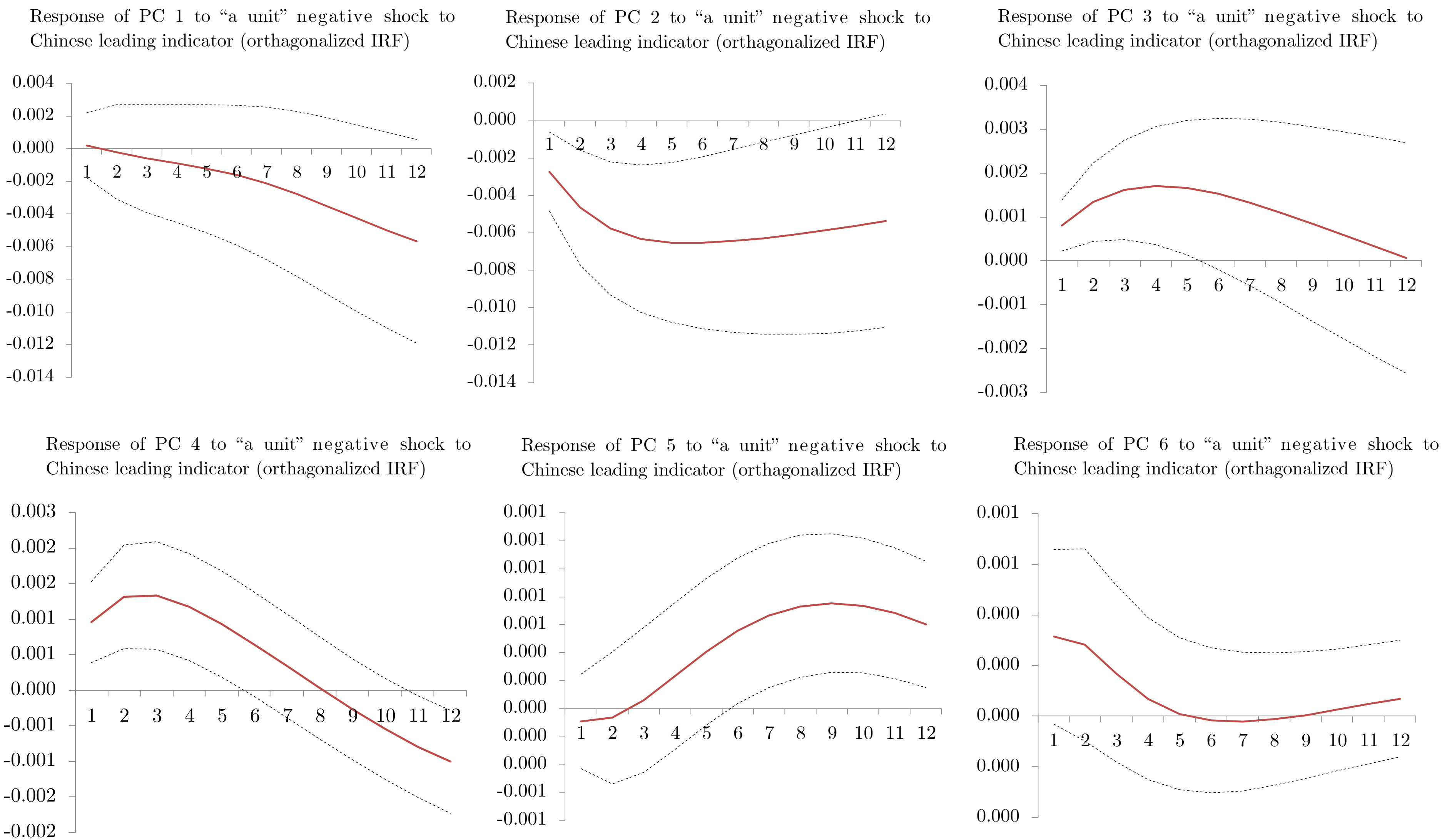


Figure 11: Orthogonalized responses of the six principal components to the negative impulse to the Chinese leading indicator (Cholesky identification scheme with lower triangular variance-covariance matrix). Sample spans from June 2009 to December 2017. Variables included in the VAR are ordered as in the second row of Table 5.

Table 6 presents the estimated prices of risks of the six-factor model for the German economy. The risk of the first principal component is affected significantly by itself, German unemployment, German leading indicator, and the Chinese leading indicator. Risks of the second principal component are affected significantly by the first, the second, the third, and the sixth principal component. The risks of the third principal component are not affected significantly in the German case. The risks of the fourth principal component are affected significantly by the first, the second, the fourth principal component, the US leading indicator, and the Chinese leading indicator. The risks of the fifth PC are affected significantly by the second principal component, by itself, the sixth principal component, German core inflation, and the German leading indicator. The sixth PC is affected significantly by the third, the fourth, and the sixth PC.

Table 6: Estimated prices of risk, λ_0^s and $\lambda_1^{s\cdot}$ of the two-country affine term structure model using an estimator as outlined in Diez de Los Rios (2018). Sample spans from June 2009 to December 2017. Spanned factors: $X_t^s = [PC\ 1_t \quad PC\ 2_t \quad PC\ 3_t \quad PC\ 4_t \quad PC\ 5_t \quad PC\ 6_t]'$. Unspanned factors: $X_t^u = [ur_{GER,t} \quad CCPI_{GER,t} \quad CLI_{GER,t} \quad CLI_{CH,t}]'$. Bolded coefficients are significant at the 10% level. I present the remaining estimated parameters in Appendix A.1. $ur_{GER,t}$ – German unemployment rate, $CCPI_{GER,t}$ – German core inflation rate, $CLI_{GER,t}$ – German leading indicator, $CLI_{CH,t}$ – Chinese leading indicator.

Factor	λ_0	$\lambda_{1,1}$	$\lambda_{1,2}$	$\lambda_{1,3}$	$\lambda_{1,4}$	$\lambda_{1,5}$	$\lambda_{1,6}$	$\lambda_{1,7}$	$\lambda_{1,8}$	$\lambda_{1,9}$	$\lambda_{1,10}$
	(constant)	(PC1)	(PC2)	(PC3)	(PC4)	(PC5)	(PC6)	(ur_{ger})	($CCPI_{ger}$)	(CLI_{ger})	(CLI_{ch})
PC 1	**-0.5327**	**-0.1591**	-0.0177	0.1200	-0.2197	-0.0184	-0.7666	**13.6070**	5.4407	**1.4448**	**-2.4243**
(t-statistic)	-3.059	-2.637	-0.430	1.041	-0.685	-0.045	-1.284	3.880	1.342	1.980	-3.452
PC 2	-0.1132	**-0.1075**	**-0.0852**	**0.2302**	0.0886	-0.0790	**-1.2165**	3.2420	0.7140	0.4637	-0.2359
(t-statistic)	-0.637	-1.745	-2.025	1.952	0.270	-0.189	-1.994	0.907	0.173	0.623	-0.329
PC 3	-0.0108	-0.0001	0.0143	-0.0209	0.1269	-0.1951	0.1786	-0.1504	-0.0249	-0.0014	-0.0387
(t-statistic)	-0.215	-0.005	1.140	-0.590	1.323	-1.573	0.966	-0.150	-0.022	-0.006	-0.193
PC 4	0.0619	**0.0336**	**-0.0263**	-0.0079	**-0.4359**	-0.1193	0.1715	-0.3215	0.3008	**0.6114**	**-0.6186**
(t-statistic)	1.330	2.064	-2.255	-0.240	-4.896	-1.033	0.993	-0.349	0.283	3.193	-3.356
PC 5	**-0.0489**	-0.0139	**0.0166**	0.0181	0.0916	**-0.2317**	**-0.2584**	0.8703	**-1.1762**	**-0.4057**	-0.0531
(t-statistic)	-1.758	-1.413	2.166	0.818	1.620	-3.079	-2.262	1.638	-1.915	-3.672	-0.499
PC 6	0.0111	0.0012	-0.0107	**0.0466**	**-0.1659**	0.0825	**-0.2270**	0.2680	-0.2612	0.1111	-0.1559
(t-statistic)	0.394	0.124	-1.572	2.421	-3.139	1.214	-2.258	0.475	-0.401	0.946	-1.380

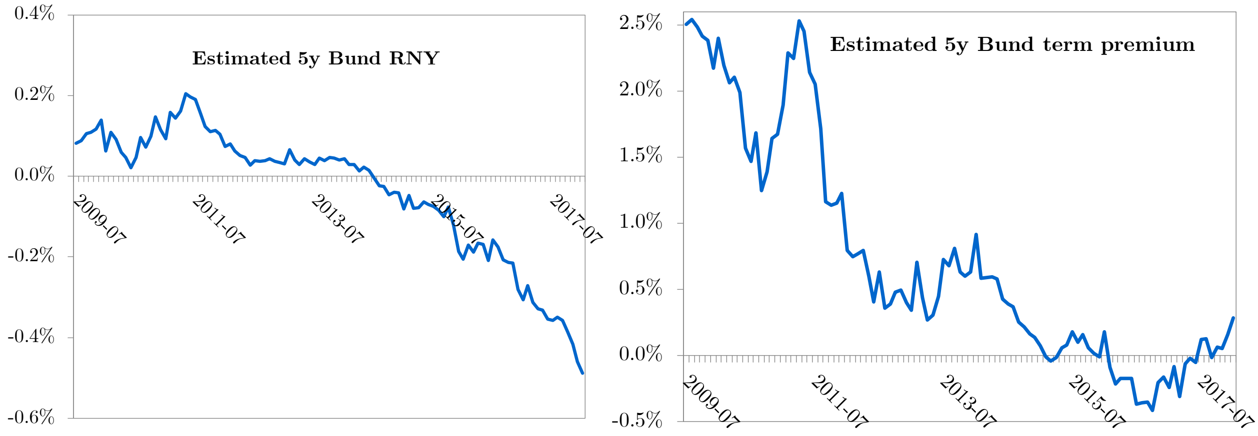


Figure 12: Model implied 5-year Bund risk-neutral yield (expected future nominal short rate) (left panel) and the 5y Bund term premium (right panel) estimated with the six-factor model (which uses PC1 to PC6) and unspanned macroeconomic variables: German unemployment, German core inflation, German leading indicator, and the Chinese leading indicator. I use an estimator as outlined in Diez de Los Rios (2018). Sample spans from June 2009 to December 2017.

The 5y Bund risk-neutral yield presented in the left panel of Figure 12 decreased from 8 basis points in June 2009 to −49 basis points by December 2017. The 5y Bund term premium, depicted in the right panel of Figure 12, decreased from 2.5 percent in June 2009 to 1 percent in August 2010. It increased back to 2.5 percent by March 2011. After the sovereign debt crisis, the 5y Bund term premium decreased to 30 basis points by March 2013. By December 2013, the 5y Bund term premium increased to 90 basis points. It decreased to −40 basis points by July 2016. The 5y Bund term premium increased to 30 basis points by December 2017.

To measure the effect of the Chinese leading indicator on the German yield curve, I perform a similar exercise as in the US case. I increase the principal components by the significant coefficients $\hat{\lambda}_{1,10}$ which are estimated in Table 6 and multiply them by −0.01. I calculate the change of the in-sample mean of the 5y Treasury yield, the 5y Treasury risk-neutral yield and the 5y Treasury term premium which are presented in Figure 12.

A *one percentage point decrease of the Chinese leading indicator increases the model implied 5y Bund yield by 3.8 basis points, the model implied 5y Bund term premium by 3.3 basis points, and the model implied 5y Bund risk-neutral by 0.5 basis points*[17]. *However, using the four-factor single country model, the model implied 5y Bund yield and its term premium decrease by 0.5 basis points in the post sovereign debt crisis sample (from December 2011 to December 2017). The 95 percent confidence interval in the post sovereign debt crisis sample includes zero effects.*

The 95 percent confidence interval (in basis points) for a percentage points decrease of the Chinese leading indicator of the model implied 5y Bund yield is $(1.7,\ 5.8)$, the model implied 5y Bund term premium is $(1.5,\ 5)$ and of the model implied 5y Bund risk-neutral yield is $(0.2,\ 0.8)$. The 95 percent confidence interval (in basis points) in the post sovereign debt crisis sample using a four-factor single-country model of the model implied 5y Bund yield is $(-2.4,\ 1.4)$, the model implied 5y Bund term premium is $(-2.43,\ 1.34)$ and of the model implied 5y Bund risk-neutral yield is $(0.03,\ 0.06)$.

Figure 13 presents the impulse response functions of the six principal components to "a unit" shock to the Chinese leading indicator in the model which includes the German macroeconomic variables. The variables which are included in the model are the six principal components (extracted jointly from the US and German yield curve), German unemployment, German core inflation, the German leading indicator, and the Chinese leading indicator.

The response of the first principal component equals -0.0041 in the 1st month. The response of the first principal component increases in absolute terms to -0.0072 in the 5th month from where it reverts back to 0. Again, I divide the responses by 0.00168 which is equal to a response of the Chinese leading indicator on itself in period zero. Afterwards, I multiply the

---

[17] When I do not condition on German unemployment, German core inflation and German leading indicator, average estimated effects in Germany decrease to (in absolute terms): the model implied 5y Bund yield decreases by -0.1 basis points, the model implied 5y Bund term premium by -0.15 basis points, and the model implied 5y Bund risk-neutral increases by 0.05 basis points.

response by -0.01 to quantify a one percentage point decrease. In Figure 13, we can observe that the responses of the fourth and the fifth principal components are significant as well.

To measure the long-run effects of the Chinese leading indicator on the German yield curve, I increase the principal components by the significant responses in the 5^{th} month when the responses of the 1^{st}, the 4^{th}, and the 5^{th} PCs are significant and the largest in absolute terms.

A one percentage point decrease of the Chinese leading indicator increases the in-sample average of the model implied 5y Bund yield by 5.9 basis points, the 5y Bund term premium by 4.7 basis points and the 5y Bund risk-neutral yield by 1.2 basis points in the 5^{th} *month. The 5y Bund yield and the 5y Bund term premium increase by 0.6 basis points in the* 12^{th} *month.*

However, using the four-factor single country model, in the 12^{th} *month, the model implied 5y Bund yield decreases by 22.5 basis points, the 5y Bund term premium by 21.9 basis points and the 5y Bund risk-neutral yield by 0.6 basis points in the post sovereign debt crisis sample (from December 2011 to December 2017).*

The 95 percent confidence interval (in basis points) for a percentage points decrease of the Chinese leading indicator of the model implied 5y Bund yield in the 5^{th} month is $(1.2,\ 10.6)$, the model implied 5y Bund term premium is $(0.7,\ 8.6)$ and of the model implied 5y Bund risk-neutral yield is $(0.5,\ 2)$.

The 95 percent confidence interval (in basis points) for a percentage points decrease of the Chinese leading indicator in the 12^{th} month in the post sovereign debt crisis sample using a single-country model of the model implied 5y Bund yield is $(-17.3,\ -27.7)$, the model implied 5y Bund term premium is $(-16.9,\ -27)$ and of the model implied 5y Bund risk-neutral yield is $(-0.4,\ -0.7)$. The impulse response functions of the first four principal components extracted from the German yield curve to a positive impulse to the Chinese leading indicator in the post sovereign debt crisis sample are depicted in Appendix A.2.

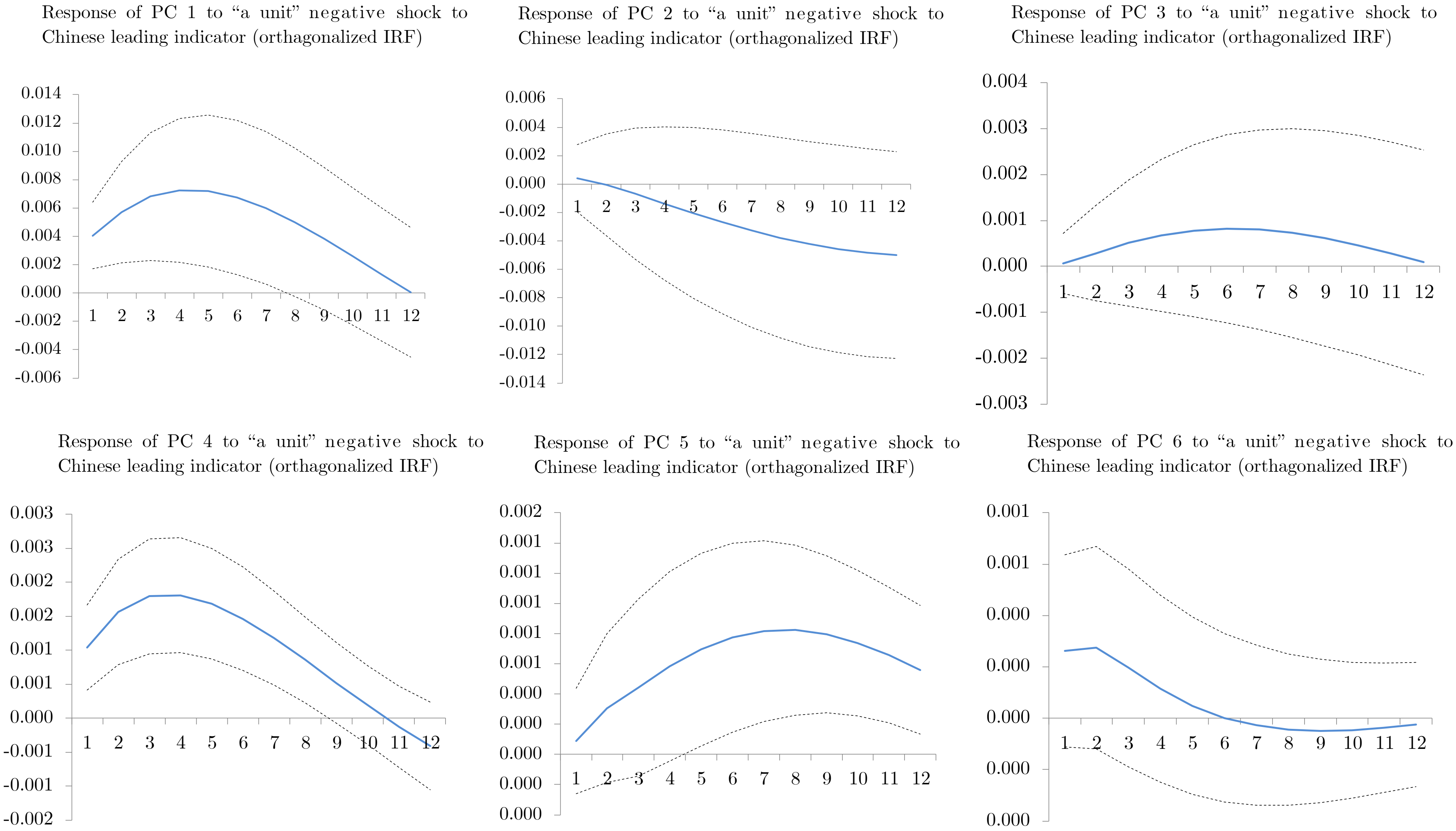


Figure 13: Orthogonalized responses of the six principal components to the negative impulse to the Chinese leading indicator (Cholesky identification scheme with lower triangular variance-covariance matrix). Sample spans from June 2009 to December 2017. Variables included in the VAR are ordered as in the second row of Table 6.

5. Conclusions

I estimate the joint affine term structure model of the US and German yield curves with the unspanned macroeconomic variables which include the Chinese leading indicator. I decompose the 5y nominal interest rates in the US and Germany in the 5y risk-neutral yields and the 5y term premia. I investigate how important is a Chinese slowdown we are observing after the financial crisis for the 5y nominal interest rates in the US and Germany, the 5y risk-neutral yields, and the 5y term premia.

I measure a Chinese slowdown as a growth differential between China and the US/Germany, which I empirically represent with the changes in the leading indicators. For each economy, in the model, I include the six principal components extracted jointly from the US and German yield curves, the domestic unemployment rate, the domestic core inflation rate, the domestic leading indicator, and the Chinese leading indicator.

A one percentage point lower Chinese leading indicator lowers the 5y Treasury yield and the 5y Treasury term premium by 4.1 basis points over the short run. In the 5^{th} month, the 5y Treasury yield decreases by 10.2 basis points, the 5y Treasury term premium by 9.2 basis points, and the 5y Treasury risk-neutral yield by 1 basis point. The 5y Bund yield *increases* by 3.8 basis points, the 5y Bund term premium by 3.3 basis points, and the 5y Bund risk-neutral yield by 0.5 basis points over the short run. In the 5^{th} month, the responses strengthen to 5.9 basis points, 4.7 basis points, and 1.2 basis points. However, the higher 5y Bund term premium could be driven by the ongoing sovereign debt crisis in the euro area.

I re-estimate the four-factor single country affine term structure model in the post sovereign debt crisis sample for the German economy. In the 12^{th} month, the model implied 5y Bund yield decreases by 22.5 basis points, the 5y Bund term premium by 21.9 basis points and the 5y Bund risk-neutral yield by 0.6 basis points.

My empirical findings suggest that the lower Chinese leading indicator helped to decrease the 5y Treasury yield and its term premium after the financial crisis, and the 5y Bund yield and its term premium after the sovereign debt crisis. Long-term bonds provide a hedge against the risks of lower growth and inflation when the monetary policy is constrained by the effective lower bound. Bondholders are willing to accept lower compensation for bearing the duration risk, the 5y term premium. In such an environment, investors became particularly sensitive towards the signals about the future growth and inflation risks such as deteriorating outlook of the Chinese economy.

References

1. Abrahams, M., Adrian, T., Crump, R. K., Moench, E., and Yu, R., 2016, Decomposing real and nominal yield curves. *Journal of Monetary Economics*, 84, 182-200.
2. Adrian, T., Crump, R. K., and Moench E., 2013, Pricing the Term Structure with Linear Regressions, *Journal of Financial Economics*, 110(1), 110-138.
3. Baele, L., Driessen, J., Ebert, S., Londono, J. M., and Spalt O. G., 2018, Cumulative prospect theory, option returns, and the variance premium, *The Review of Financial Studies*, 32(9), 3667-3723.
4. Bauer, M. D., and Rudebusch, G. D., 2014, The Signaling Channel for Federal Reserve Bond Purchases, *International Journal of Central Banking.*
5. Bernanke, B. S., 2015, Why are interest rates so low, part 4: Term premiums, *Brookings blog post.*
6. Campbell, J. Y., Sunderam, A., and Viceira L.M., 2009, Inflation bets or deflation hedges? The changing risks of nominal bonds, *Working paper*, National Bureau of Economic Research.
7. Cashin, P., Mohaddes, K., and Raissi, M., 2017, China's slowdown and global financial market volatility: Is world growth losing out?, *Emerging Markets Review*, 31, 164-175.
8. Christensen, J. H., Lopez, J. A., and Rudebusch, G. D., 2010, Inflation expectations and risk premiums in an arbitrage-free model of nominal and real bond yields, *Journal of Money, Credit and Banking*, 42, 143-178.
9. Cochrane, J. H., and Piazzesi, M., 2005, Bond risk premia, *American Economic Review*, 95(1), 138-160.
10. Diez de Los Rios, A., 2015, A new linear estimator for Gaussian dynamic term structure models, *Journal of Business & Economic Statistics*, 33(2), 282-295.
11. Diez de Los Rios, A., 2018, Optimal Estimation of Multi-Country Gaussian Dynamic Term Structure Models Using Linear Regressions, *Working paper.*
12. European Central Bank, ECB, 2017, ECB Economic Bulletin. *Frankfurt am Main,* 07/2017.

13. Gauvin, L., and Rebillard, C. C., 2015, Towards recoupling? Assessing the global impact of a Chinese hard landing through trade and commodity price channels, *The World Economy*.
14. Gürkaynak, R. S., B. Sack, and J.H. Wright, 2007, The US Treasury yield curve: 1961 to the Present, *Journal of Monetary Economics*, 54(8), 2291-2304.
15. Hördahl, P., and Tristani, O., 2012, Inflation risk premia in the term structure of interest rates. *Journal of the European Economic Association*, 10(3), 634-657.
16. Maletic, M., 2018, Chinese foreign reserves and the US yield curve, *Working paper*.
17. Metelli, L., and Natoli, F., 2017, The effect of a Chinese slowdown on inflation in the euro area and the United States, *Economic Modelling*, 62, 16-22.
18. Roussellet, G., 2017, Affine Term Structure Modeling and Macroeconomic Risks at the Zero Lower Bound, *Working paper*.

A.1. Estimated parameters of the joint affine model of the US and German nominal term structures in the post financial crisis sample

$\delta_0^{US,s}$	$\delta_0^{GER,s}$
(constant)	(constant)
-0.0062	**-0.0035**
0.0011	0.0019
(standard error)	(standard error)

Factor	$\delta_1^{US,s}$	$\delta_1^{GER,s}$
PC 1	**-0.0333**	**0.0729**
(standard error)	0.0005	0.0009
PC 2	**0.0549**	**-0.0478**
(standard error)	0.0010	0.0018
PC 3	**0.1900**	**-0.0444**
(standard error)	0.0033	0.0057
PC 4	**0.1676**	**0.2083**
(standard error)	0.0059	0.0101
PC 5	**-0.2907**	**-0.1817**
(standard error)	0.0100	0.0172
PC 6	**-0.2495**	**0.5342**
(standard error)	0.0163	0.0280

Factor	$\mu_s^\star$
	(constant)
PC 1	**0.0016**
(standard error)	0.0003
PC 2	**0.0033**
(standard error)	0.0004
PC 3	**0.0032**
(standard error)	0.0009
PC 4	**0.0034**
(standard error)	0.0009
PC 5	**-0.0049**
(standard error)	0.0015
PC 6	**-0.0012**
(standard error)	0.0003

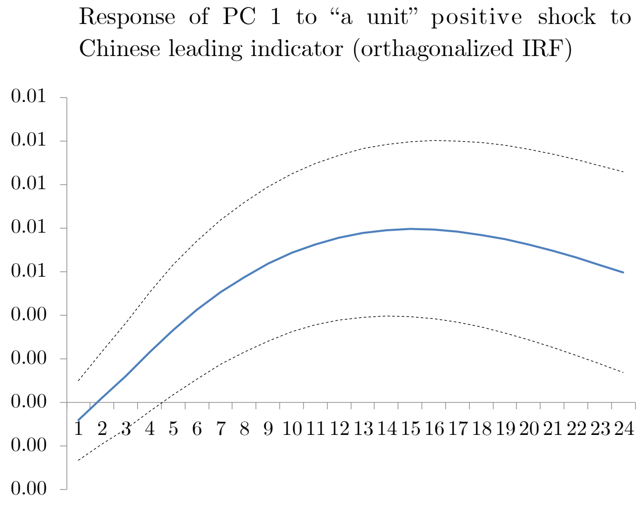
Factor	$\Phi_{ss,1,1}^\star$	$\Phi_{ss,1,2}^\star$	$\Phi_{ss,1,3}^\star$	$\Phi_{ss,1,4}^\star$	$\Phi_{ss,1,5}^\star$	$\Phi_{ss,1,6}^\star$
	(PC1)	(PC2)	(PC3)	(PC4)	(PC5)	(PC6)
PC 1	**1.0239**	**0.0168**	**0.0270**	**-0.0697**	**0.0156**	**-0.2475**
(standard error)	0.0001	0.0003	0.0009	0.0016	0.0028	0.0044
PC 2	**0.0132**	**1.0274**	**-0.0804**	**-0.0588**	**0.0619**	**0.1536**
(standard error)	0.0002	0.0004	0.0014	0.0023	0.0039	0.0061
PC 3	0.0003	**0.0194**	**1.0031**	**-0.0856**	**0.1813**	**0.2084**
(standard error)	0.0004	0.0008	0.0028	0.0048	0.0080	0.0157
PC 4	**0.0127**	**0.0227**	**-0.0083**	**0.9695**	**0.2218**	**-0.2932**
(standard error)	0.0005	0.0009	0.0030	0.0052	0.0087	0.0169
PC 5	**0.0021**	0.0010	**-0.0108**	0.0015	**0.9408**	0.0397
(standard error)	0.0007	0.0014	0.0047	0.0080	0.0135	0.0241
PC 6	**0.0030**	**0.0041**	**-0.0397**	**0.0580**	**0.0114**	**0.6562**
(standard error)	0.0001	0.0003	0.0009	0.0015	0.0025	0.0052

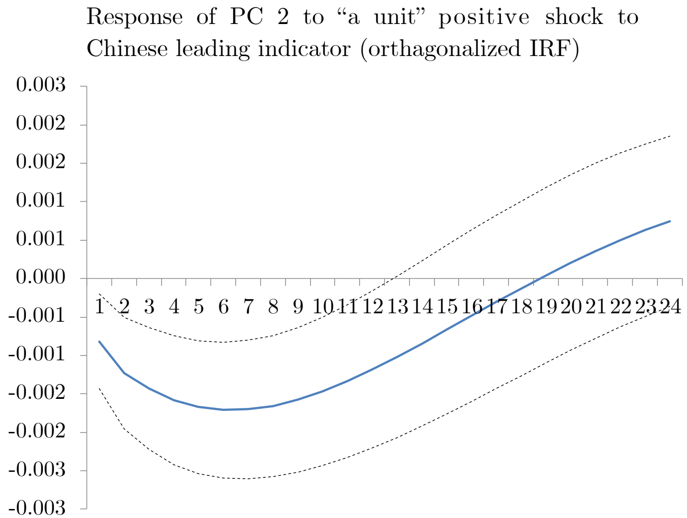
	Eigenvalues
λ_1	0.7491 + 0.0000i
λ_2	0.9417 + 0.0409i
λ_3	0.9417 - 0.0409i
λ_4	1.0009 + 0.0249i
λ_5	1.0009 - 0.0249i
λ_6	0.9865 + 0.0000i

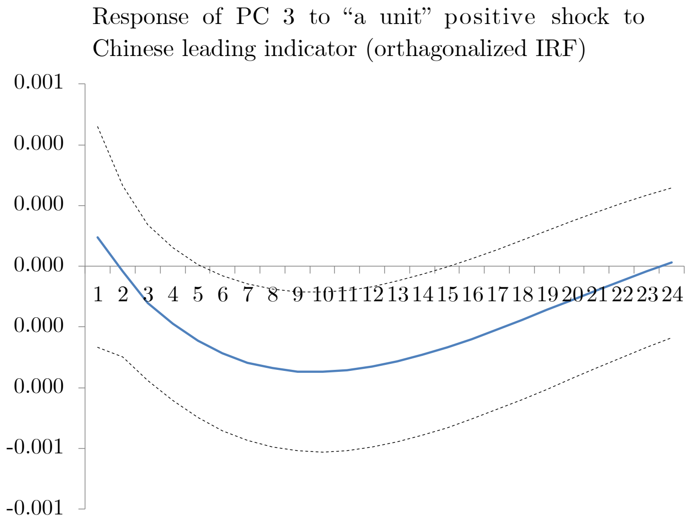
Factor	$\Sigma_{1,1}^{Chol}$	$\Sigma_{1,2}^{Chol}$	$\Sigma_{1,3}^{Chol}$	$\Sigma_{1,4}^{Chol}$	$\Sigma_{1,5}^{Chol}$	$\Sigma_{1,6}^{Chol}$	$\Sigma_{1,7}^{Chol}$	$\Sigma_{1,8}^{Chol}$	$\Sigma_{1,9}^{Chol}$	$\Sigma_{1,10}^{Chol}$
	(PC1)	(PC2)	(PC3)	(PC4)	(PC5)	(PC6)	(ur_{us})	($CCPI_{us}$)	(CLI_{us})	(CLI_{ch})
PC 1	**0.0781**									
(standard error)	0.0058									
PC 2	**0.0272**	**0.0749**								
(standard error)	0.0087	0.0044								
PC 3	0.0023	**-0.0124**	**0.0181**							
(standard error)	0.0020	0.0019	0.0015							
PC 4	**0.0069**	-0.0031	**0.0054**	**0.0192**						
(standard error)	0.0021	0.0020	0.0024	0.0013						
PC 5	**0.0023**	-0.0007	0.0011	-0.0002	**0.0126**					
(standard error)	0.0013	0.0012	0.0014	0.0013	0.0009					
PC 6	-0.0001	**0.0031**	0.0000	**0.0057**	0.0001	**0.0114**				
(standard error)	0.0012	0.0011	0.0013	0.0013	0.0012	0.0008				
ur_{us}	0.0002	**-0.0004**	-0.0001	0.0000	**0.0003**	**0.0003**	**0.0013**			
(standard error)	0.0001	0.0001	0.0002	0.0001	0.0001	0.0001	0.0001			
$CCPI_{us}$	-0.0001	0.0000	0.0001	-0.0001	0.0001	**-0.0002**	-0.0001	**0.0008**		
(standard error)	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001		
CLI_{us}	0.0002	**0.0004**	-0.0002	0.0001	-0.0002	**-0.0004**	-0.0002	-0.0002	**0.0015**	
(standard error)	0.0002	0.0002	0.0002	0.0002	0.0002	0.0002	0.0002	0.0002	0.0002	
CLI_{ch}	**0.0009**	0.0000	**-0.0006**	0.0004	0.0000	-0.0002	0.0003	-0.0003	**0.0013**	**0.0021**
(standard error)	0.0002	0.0002	0.0003	0.0002	0.0002	0.0002	0.0003	0.0003	0.0003	0.0001

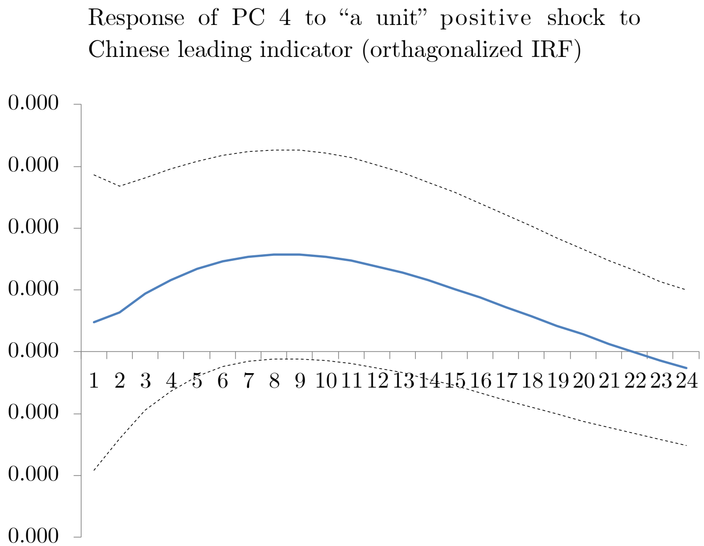
Factor	$\Sigma_{1,1}^{Chol}$	$\Sigma_{1,2}^{Chol}$	$\Sigma_{1,3}^{Chol}$	$\Sigma_{1,4}^{Chol}$	$\Sigma_{1,5}^{Chol}$	$\Sigma_{1,6}^{Chol}$	$\Sigma_{1,7}^{Chol}$	$\Sigma_{1,8}^{Chol}$	$\Sigma_{1,9}^{Chol}$	$\Sigma_{1,10}^{Chol}$
	(PC1)	(PC2)	(PC3)	(PC4)	(PC5)	(PC6)	(ur_{ger})	($CCPI_{ger}$)	(CLI_{ger})	(CLI_{ch})
PC 1	**0.0823**									
(standard error)	0.0061									
PC 2	**0.0292**	**0.0786**								
(standard error)	0.0092	0.0046								
PC 3	0.0026	**-0.0143**	**0.0184**							
(standard error)	0.0020	0.0019	0.0016							
PC 4	**0.0057**	-0.0031	**0.0052**	**0.0199**						
(standard error)	0.0022	0.0021	0.0025	0.0013						
PC 5	0.0024	-0.0001	0.0007	0.0002	**0.0122**					
(standard error)	0.0013	0.0012	0.0014	0.0013	0.0009					
PC 6	-0.0008	**0.0027**	0.0005	**0.0060**	-0.0009	**0.0114**				
(standard error)	0.0012	0.0011	0.0013	0.0013	0.0012	0.0008				
ur_{us}	0.0001	0.0000	-0.0001	0.0000	0.0000	0.0002	**0.0006**			
(standard error)	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001			
$CCPI_{us}$	0.0003	0.0001	-0.0002	**0.0004**	0.0001	0.0000	0.0001	**0.0018**		
(standard error)	0.0002	0.0002	0.0002	0.0002	0.0002	0.0002	0.0004	0.0001		
CLI_{us}	0.0004	**0.0007**	-0.0005	0.0000	0.0001	-0.0002	-0.0002	**0.0008**	**0.0022**	
(standard error)	0.0003	0.0003	0.0003	0.0003	0.0003	0.0003	0.0008	0.0004	0.0002	
CLI_{ch}	**0.0011**	**0.0006**	**-0.0009**	0.0002	0.0002	**-0.0006**	**0.0009**	**0.0007**	**0.0020**	**0.0019**
(standard error)	0.0002	0.0002	0.0003	0.0002	0.0002	0.0002	0.0005	0.0002	0.0002	0.0002

A.2. Orthogonalized responses of the four principal components extracted from the German yield curve to the positive impulse to the Chinese leading indicator (Cholesky identification scheme with lower triangular variance-covariance matrix). Sample spans from December 2011 to December 2017. Variables included in the VAR (1) model: the four principal components extracted from the German yield curve, German unemployment rate, German core inflation rate, the German leading indicator, and the Chinese leading indicator.









Chapter 2: Chinese foreign reserves and the US yield curve

Matjaz Maletic[1]

This version: 24th of October 2019

Abstract

I estimate an affine term structure model with unspanned macroeconomic variables which include the Chinese foreign reserves. The low growth of the Chinese foreign reserves after the financial crisis is more important for movements of the 5y Treasury yield and its term premium than the high growth before. It signals lower 5y Treasury yield and decreases the compensation for bearing the duration risk (the 5y Treasury term premium). The economically important feedbacks from the level factor of the US yield curve suggest that the effects are running in both directions.

1. Introduction

This paper investigates to what extent the increasing prevalence of the Chinese economy in the global economy, and the accumulation of the foreign reserves, in particular, affected the US yield curve. In 1999, the share of the Chinese economy in the global economy equaled 3.4 percent. The share of the Chinese economy increased to 15.1 percent in 2017. China has become one of the major powers in the global economy.

From 1998 to 2007, China has experienced record high growth rates, 10 percent on average. However, China was managing the exchange rate. The Renminbi was pegged to the US Dollar since 1994. In 2005, the People's Bank of China (PBOC) announced that China was moving towards a managed peg. The Renminbi preserved a tight link to the US Dollar. Specifically, the daily central parity rate was announced around which the Renminbi-US Dollar par fluctuated. The PBOC used this mechanism to hold back the appreciation of the Renminbi[2]. The foreign exchange policy has supported the export-oriented growth model of the Chinese economy. From July 2005 until December 2007 the Renminbi appreciated by 2 percent relative to its trading partners[3]. To depreciate the Renminbi, among other measures, the PBOC buys US Dollars and sells the Renminbi. The PBOC's efforts to hold back the

---

[1] Tilburg University, Finance Department. E-mail: maletic.matjaz@gmail.com

[2] Das (2019) discusses the evolution of the Chinese foreign exchange rate policy over time into the greater detail.

[3] Chinese effective exchange rate is available at http://data.imf.org/regular.aspx?key=61545850

appreciation of the Renminbi combined with the Chinese current account surplus increased the Chinese foreign exchange reserves[4]. From December 2000 to December 2007, the PBOC accumulated 1.4 trillion US dollars of foreign reserves. Figure 1 (Panel A) illustrates the different channels which were at play before the financial crisis.

After the financial crisis, China has slowed down. The average yearly growth of the real output decreased to 8.1 percent. The Renminbi continued to appreciate against its trading partners. The Chinese foreign exchange reserves increased to 4 trillion US dollars by June 2014. From December 20007 to July 2015 the Renminbi appreciated by 36 percent. The strong Renminbi represented an additional anchor in times when the economic growth was slowing down. Given a substantial appreciation of the Renminbi in 2014 due to a strong US Dollar, and economic slowdown in China, market consensus was building that the Renminbi became overvalued[5].

In July 2015, the PBOC announced a change in the foreign exchange regime and a move closer towards the market determination of the Renminbi exchange rate. Market participants interpreted the regime change as the beginning of a sizeable Renminbi depreciation, a trend reversal. The PBOC intervened in foreign exchange markets to stabilize the Renminbi. In early 2016, the PBOC published a basket of currencies it plans to follow and put the new central parity mechanism in place (Das, 2019)[6]. By December 2017, the Chinese foreign reserves decreased to 3.1 trillion US Dollars. The average yearly growth of the Chinese foreign exchange reserves decreased from 35.5 percent before the financial crisis to 10 percent after the financial crisis. Figure 1 (Panel B) illustrates the different channels which are at play after the financial crisis[7].

---

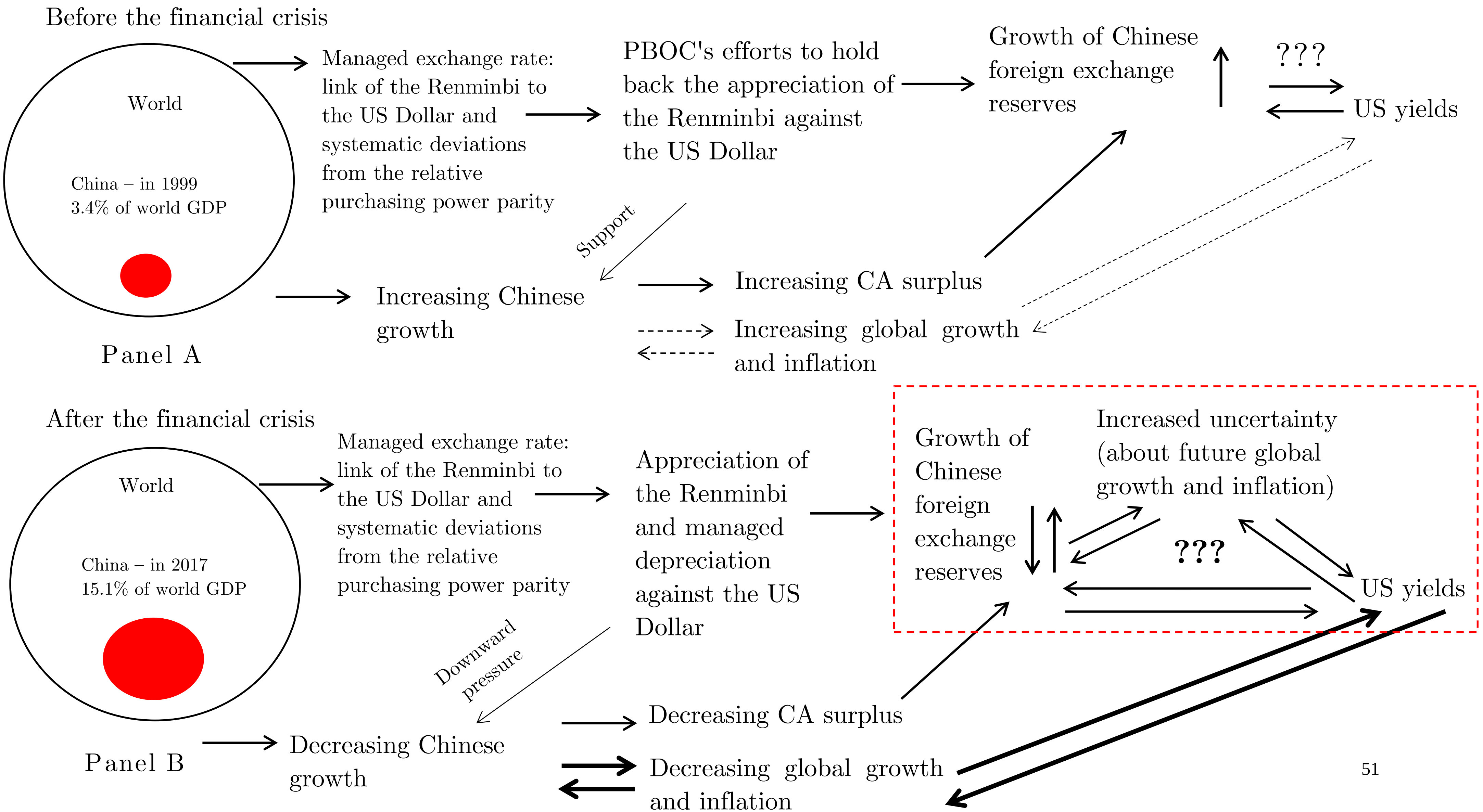
[4] The average Chinese current account surplus in period from 1999 to 2007 equaled 4.7 percent of GDP.

[5] After substantial real appreciation in 2014 IMF's staff declared that Chinese currency is no longer undervalued. The announcement is available at https://www.imf.org/en/News/Articles/2015/09/14/01/49/pr15237

[6] Basket is available at http://test.chinamoney.com.cn/english/svcnrl/20161229/2047.html

[7] The average Chinese current account surplus decreased to 2.5 percent after the financial crisis.

Figure 1: Share of the Chinese economy in the global economy (in nominal GDP), and the connection between the Chinese foreign exchange reserves, and the US yields before and after the financial crisis



My main contribution is to quantify the effects of the Chinese foreign exchange reserves on the 5y Treasury yield, the 5y Treasury term premium, and the 5y Treasury risk-neutral yield. The 5y Treasury yield can be decomposed in (1) the future expectations of the short rate (the 5y Treasury risk-neutral yield), and (2) the compensation for bearing the term (duration) risk (the 5y Treasury term premium).

I hypothesize that the Chinese foreign exchange reserves can change the 5y Treasury yield and its term premium in two opposite ways.

First, the *higher* growth of the Chinese foreign exchange reserves typically also increases the dollar amount of the US Treasuries held by the PBOC. The higher debt holdings increase the prices of the US Treasuries which lowers the US nominal interest rates. These are direct effects of the Chinese foreign reserves on the US yield curve. Such an example is presented in the right part of Panel A in Figure 1. The major part of the PBOC's holdings of the US Treasuries has a maturity of 5 years or less[8]. Therefore, the increased Chinese foreign reserves lower the 5y Treasury yield by lowering the compensation for bearing the duration risk (the 5y Treasury term premium).

Beltran, Kretchmer, Marquez and Thomas (2013), for instance, estimate the effect of an increase in foreign official holdings of US Treasuries (from all foreign countries, not just from China) on the 5y Treasury term premium, and find the effect to be significantly negative[9]. Warnock and Warnock (2009) estimate that in the period from 1984 to 2005 the 10y Treasury yield would be 80 basis points higher if there would be no foreign inflows into the US Treasuries. However, as such direct link of the foreign reserves on the US yield curve should manifest itself instantaneously the methodology developed in this paper by construction takes such direct effects out. Namely, in my empirical model, I assume that the Chinese foreign exchange reserves affect the US yield curve with a one month lag.

---

[8] In June 2017 approximately 75 percent of the total Treasury and Agency debt held by the foreign Central Banks had the maturity of 5 years or less (Department of the Treasury, 2018, Exhibit 15).

[9] The effect continues being negative even after taking into account the reaction by foreign private investors to the yield changes induced by the shocks to foreign official inflows.

Second, indirectly, the *lower* and even negative growth of the Chinese foreign reserves can be a result of (unexpected) significant depreciation of the Renminbi against the US Dollar and increased (unobserved) uncertainty about the future growth of the Chinese economy. The lower growth of Chinese foreign reserves can be a signal for *risks* related to lower future global growth and inflation. Such (unobserved) *risks* might be driving the predictive power of the Chinese foreign reserves on the US yield curve. China has become the second biggest economy in the world, and an important factor in global growth and inflation (Gauvin and Rebillard, 2015). In an environment where the monetary policy is constrained by the effective lower bound, these *risks* increase the nominal bond prices by signaling that the FED will have to keep rates lower for longer[10]. Such an example is presented in the upper right part of Panel B in Figure 1. Bondholders are willing to accept *lower* compensation for bearing the duration risk, the 5y Treasury term premium.

In my companion paper, Maletic (2018), I estimate that in the post financial crisis sample a lower Chinese leading indicator lowers the 5y Treasury yield and decreases the compensation for bearing the duration risk (the 5y Treasury term premium). In this paper, I test whether the Chinese foreign exchange reserves represent incremental information to the Chinese leading indicator for the 5y Treasury yield and its term premium. Chinese leading indicator is measuring the growth rate of the Chinese economy. It does not, however, take into account the tight link between the US Dollar and the Renminbi which is crucial for manifestation of risks, such as growth and inflation risks, in the US yield curve through the Chinese foreign exchange reserves.

The managed exchange rate forced the PBOC to follow the US Dollar by intervening in the foreign exchange market. If China would not be managing its exchange rate, the Chinese foreign reserves, which are an instrument of the foreign exchange policy, would not exhibit such a high correlation with the US Dollar. In my sample, the correlation between the yearly growth of the Dollar index and the growth of the Chinese foreign reserves is equal to -0.66.

---

[10] This interpretation follows Bernanke (2015).

I estimate the affine term structure model with unspanned macroeconomic variables in the period from December 2000 to December 2017 following Diez de Los Rios (2018). I follow Beltran, Kretchmer, Marquez and Thomas (2013) and focus on the US term structure of up to 5 years. In the model, I include the four principal components extracted from the US yield curve, the US unemployment rate, US CPI inflation, the Chinese foreign exchange reserves, Chinese CPI inflation, the Dollar index, and the Chinese leading indicator.

I motivate the inclusion of the US unemployment rate in the model because of two reasons: (1) to condition on the omitted factors such as the global output gap, and (2) to impose the Taylor rule in the short rate equation.

The efforts to limit the appreciation of the Renminbi before the financial crisis took its toll. By following the monetary policy of the country to which the Renminbi was pegged, namely the US, the PBOC risked overheating the economy[11]. Chinese inflation increased from 0 percent in 2000 to 8.5 percent in 2008[12]. Chinese inflation was unusually high in 2007 and 2008 followed by deflation in 2009.

It is important to condition on the Chinese CPI inflation and the US Dollar because a link of the Renminbi to the US Dollar drives systematic deviations of the actual exchange rate from the exchange rate implied by the relative purchasing power parity[13]. Such deviations could be driving risks related to the future growth and inflation of the Chinese economy, and increase the predictive information of the Chinese foreign reserves for the US yield curve. Bini Smaghi (2010) argues that anchoring the exchange rate involves several medium-term distortions, cost, and risks. The pickup of inflation rates in Emerging Market Economies, and China, in

---

[11] Among others, Bini Smaghi (2010) stresses out that by managing their exchange rates emerging market economies imported the monetary policy stances of the developed economies which became suboptimal for the monetary policy stance of the emerging market economy. Bernanke (2017) discusses the US monetary policy in international context.

[12] Kroeber (2011) points out that by keeping the Renminbi undervalued for too long PBOC could risk increasing the inflation so high that it could begin harming the Chinese real economic activity. Kroeber (2011) and Frankel (2015) discuss political aspects of Chinese monetary and exchange rate policies.

[13] Relative purchasing power parity implies that the inflation differential should be reflected in the expected spot exchange rate.

particular, is one of them. In particular, the higher Chinese inflation relative to the US drives the expected depreciation of the Renminbi against the US Dollar (assuming the real exchange rate remains the same).

To test the incremental information of the Chinese foreign reserves, the Chinese leading indicator should be included in the model as well. The principal components control for the feedback-loop from the US yield curve to the macroeconomic variables.

My main empirical results yield several new findings.

A 10 percentage point negative shock to the Chinese foreign reserves decreases the model implied 5y Treasury yield by 5.1 basis points over the short run. The 5y Treasury term premium decreases by 4.9 basis points and the 5y Treasury risk-neutral yield by 0.2 basis points. In the 24th month, the model implied 5y Treasury yield decreases by 34.9 basis points, the model implied 5y Treasury term premium by 28.8 basis points and the 5y Treasury risk-neutral yield by 6.1 basis points. In the 24th month, the model implied 5y Treasury yield decreases by 28.2 basis points, the model implied 5y Treasury term premium by 21.9 basis points and the 5y Treasury risk-neutral yield by 6.3 basis points, when in addition I control for the nominal Renminbi effective exchange rate, the exchange rate of the Renminbi against the US Dollar, and the VIX index[14].

In 2017, China and Japan were the biggest foreign holders of the US Treasuries. China, however, is managing its exchange rate against the US Dollar. Changes in Chinese foreign reserves became important for understanding the lower 5y Treasury yield, and especially unusually low 5y Treasury term premium we are observing after the financial crisis. The lower growth of the Chinese foreign reserves signals a lower 5y Treasury yield and decreases the compensation for bearing the duration risk (the 5y Treasury term premium). The (unobserved) risks related to the (unexpected) significant depreciation of the Renminbi against the US Dollar, and the lower future growth of the Chinese economy, are decreasing

---

[14] VIX measures market expectation of near term volatility conveyed by stock index option prices. Available at https://fred.stlouisfed.org/series/VIXCLS

the Chinese foreign reserves, the 5y Treasury yield, and its term premium. China has become the second biggest economy in the world, an important player in global commodity markets, and hence an important contributor to global growth and inflation.

My empirical findings suggest that the lower growth of the Chinese foreign reserves after the financial crisis is more important for the 5y Treasury yield and its term premium than a relatively high growth of the Chinese foreign reserves before the financial crisis.

However, Chinese debt holdings are akin to feedback-loop when considering the US nominal interest rates. One percentage point increase of the 5y Treasury yield *increases* the growth of the Chinese foreign reserves by 80 basis points. In the 24th month, the response of the Chinese foreign reserves is still significant and equals 4.75 *percentage points*[15]. The positive direction is driven by the level factor of the US yield curve. Assuming that the (uncovered) interest rate parity holds higher US nominal interest rates depreciate the US Dollar against the US trading partners in the future[16]. As depreciating US Dollar tends to be accompanied by the accumulation of the Chinese foreign reserves, higher US nominal interest rates push the Chinese foreign reserves up as well. This economic mechanism helps to explain a relatively important contribution of the level factor for movements in the 5y Treasury term premium. Over the short run, a unit increase of the level factor increases the 5y Treasury yield by 13.2 basis points, the 5y Treasury term premium by 11 basis points and the 5y Treasury risk-neutral yield by 2.3 basis points.

The Chinese official sector decreased the accumulation of foreign reserves after the financial crisis. Move towards the increased flexibility of the Renminbi exchange rate to address the significant appreciation of the Renminbi in late 2014 and early 2015 forced the Chinese

---

[15] While the effect seems to be high at the first glance, the Chinese foreign reserves increased by 18x in the period from December 2000 to December 2017. The 5y Treasury yield decreased by 3 percentage points in the same period. Assuming duration of the Chinese foreign reserves equal to 10 and increase in the US nominal short rate which corresponds 1 to 1 to the 5y Treasury yield, the Chinese foreign reserves should *decrease* by 1.3x due to the lower US nominal short rate.

[16] Interest rate parity follows logic that assets with the same risk should yield the same rate of return in Renminbi and US Dollars.

official sector to intervene in the foreign exchange markets to prevent significant depreciation of the Renminbi in late 2015. Such interventions have lowered the growth of the Chinese foreign exchange reserves after the financial crisis (Das, 2019).

The lower growth of the Chinese foreign reserves provides a signal for lower US long-term interest rates in the future. In the US after the financial crisis, we are observing low levels of growth and inflation with the monetary policy constrained at the effective lower bound. In such an environment, the nominal bonds provide a hedge against the risks of lower growth and inflation. Bondholders are willing to accept lower compensation for holding long-term nominal bonds instead of short-term securities. This has pushed down the 5y Treasury term premium. My empirical findings suggest that the lower growth of the Chinese foreign reserves provides economically important signals about the lower term premium of the 5y Treasuries after conditioning on the US Dollar, the Renminbi effective exchange rate, the exchange rate of Renminbi against the US Dollar, US unemployment rate and inflation, the growth and inflation rate of the Chinese economy, and the VIX index. Feedbacks from the level factor of the US yield curve on the growth of the Chinese foreign reserves suggest that the effects are running in both directions.

The rest of this paper is organized as follows. Section 2 introduces an affine term structure model. Section 3 presents the data. Main results are presented in Section 4. Section 5 concludes.

2. Affine Term Structure Model

I follow Diez de Los Rios (2015 and 2018). His estimator has a limiting distribution which is asymptotically equivalent to the maximum likelihood estimation. The estimation starts with the joint vector-autoregressive (VAR) process for spanned and unspanned factors under the historical measure

$$\begin{bmatrix} X_t^s \\ X_t^u \end{bmatrix} = \mu + \Phi \begin{bmatrix} X_{t-1}^s \\ X_{t-1}^u \end{bmatrix} + \begin{bmatrix} v_t^s \\ v_t^u \end{bmatrix} \tag{1}$$

Where

X_t^s – spanned pricing factors (principal components) $\in \mathcal{R}^{K_s \times 1}$

X_t^u – unspanned macroeconomic variables $\in \mathcal{R}^{K_u \times 1}$

Shocks, $v_t = [v_t^s \quad v_t^u]'$, conditionally on lagged principal components and unspanned macroeconomic variables follow a Normal distribution, $v_t | \{X_s\}_{s=0}^{t-1} \sim N(0, \Sigma)$. μ, Φ, and Σ are partitioned according to the spanned factors and unspanned macroeconomic variables. Namely,

$$\mu = \begin{bmatrix} \mu_s \\ \mu_u \end{bmatrix}, \qquad \Phi = \begin{bmatrix} \Phi_{ss} & \Phi_{su} \\ \Phi_{us} & \Phi_{uu} \end{bmatrix}, \qquad and\ \Sigma = \begin{bmatrix} \Sigma_{ss} & \Sigma_{su} \\ \Sigma_{us} & \Sigma_{uu} \end{bmatrix}. \tag{2}$$

I use the principal component analysis and extract the principal components from the US yield curve which affect the bond prices directly (X_t^s). The macroeconomic variables (X_t^u) affect the bond prices merely through the principal components with a lag. In the model, I include the US unemployment rate, US CPI inflation, Chinese foreign reserves, Chinese CPI inflation, the Chinese leading indicator, and the US Dollar index.

The bond pricing factors (principal components) and the nominal short-term interest rate are related through the affine relation

$$r_t = \delta_0^s + \delta_1^{s\prime} X_t^s \tag{3}$$

Under the risk-neutral probability measure, the spanned and unspanned factors follow the VAR (1) process

$$\begin{bmatrix} X_t^s \\ X_t^u \end{bmatrix} = \begin{bmatrix} \mu_s^\star \\ \mu_u^\star \end{bmatrix} + \begin{bmatrix} \Phi_{ss}^\star & 0 \\ \Phi_{us}^\star & \Phi_{uu}^\star \end{bmatrix} \begin{bmatrix} X_{t-1}^s \\ X_{t-1}^u \end{bmatrix} + \begin{bmatrix} \upsilon_t^{s\star} \\ \upsilon_t^{u\star} \end{bmatrix} \tag{4}$$

Because unspanned macro factors do not affect bond prices under the risk-neutral measure the pricing (risk-neutral) transition matrices, $\mu^\star$ and $\Phi^\star$, can be written as

$$\mu^\star = \begin{bmatrix} \mu_s - \lambda_0^s \\ \mu_u - \lambda_0^u \end{bmatrix} = \begin{bmatrix} \mu_s - \lambda_0^s \\ \mu_u \end{bmatrix} = \begin{bmatrix} \mu_s^\star \\ \mu_u^\star \end{bmatrix}, \qquad \Phi^\star = \begin{bmatrix} \Phi_{ss} - \lambda_1^{ss} & \Phi_{su} - \lambda_1^{su} \\ \Phi_{us} - \lambda_1^{us} & \Phi_{uu} - \lambda_1^{uu} \end{bmatrix} = \begin{bmatrix} \Phi_{ss} - \lambda_1^{ss} & 0 \\ \Phi_{us} & \Phi_{uu} \end{bmatrix} = \begin{bmatrix} \Phi_{ss}^\star & 0 \\ \Phi_{us}^\star & \Phi_{uu}^\star \end{bmatrix} \tag{5}$$

Therefore, $\lambda_0^u = 0$, $\lambda_1^{us} = 0 \in \mathcal{R}^{K_u \times K_s}$, $\lambda_1^{uu} = 0 \in \mathcal{R}^{K_u \times K_u}$, the upper right $K_s \times K_u$ block of risk-neutral matrix $\Phi^\star$, $\Phi_{su}^\star = (\Phi_{su} - \lambda_1^{su})$ is zero, and $\boldsymbol{\Phi_{su} = \lambda_1^{su}} \in \mathcal{R}^{K_s \times K_u}$. Shocks, $\upsilon_t{}^\star = [\upsilon_t^{s\star} \quad \upsilon_t^{u\star}]'$, conditionally on lagged principal components and unspanned macroeconomic variables follow a Normal distribution, $\upsilon_t{}^\star | \{X_s\}_{s=0}^{t-1} \sim N(0, \Sigma)$. Σ is the same matrix as in (1).

Given the assumptions (1) – (5) bond prices are exponentially affine in the spanned factors

$$lnP_t^{(n)} = A_n^s + B_n^{s\prime} X_t^s, \tag{6}$$

The continuously compounded yield on a n-period zero-coupon bond at time t equals $y_t^{(n)} = -\frac{1}{n} lnP_t^{(n)}$, and can be written as

$$y_t^{(n)} = a_n^s + b_n^{s\prime} X_t^s, \tag{7}$$

where $a_n^s = -\frac{A_n^s}{n}$ and $b_n^s = -\frac{B_n^s}{n}$.

Following Diez de Los Rios (2018) recursive linear restrictions A_n^s and $B_n^{s\prime}$ are given as $(for\; n > 0)$

$$A_n^s = A_{n-1}^s + B_{n-1}^s{}'(\mu_s - \lambda_0^s) + \frac{1}{2} B_{n-1}^s{}' \Sigma_{ss} B_{n-1}^s - \delta_0^s \tag{8}$$

$$B_n^{s\prime} = B_{n-1}^s{}'(\Phi_{ss} - \lambda_1^{ss}) - \delta_1^{s\prime} \tag{9}$$

$$A_0^s = 0, \qquad A_1^s = -\delta_0^s, \qquad B_0^{s\prime} = 0, \qquad B_1^{s\prime} = -\delta_1^{s\prime}. \tag{10}$$

When the prices of the risk parameters λ_0^s and λ_1^{ss} in (8) and (9) are set to zero, the recursions generate the risk adjusted bond pricing parameters

$$A_n^{s,RF} = A_{n-1}^{s,RF} + B_{n-1}^{s,RF'}\mu_s + \frac{1}{2}B_{n-1}^{s,RF'}\Sigma_{ss}B_{n-1}^{s,RF} - \delta_0^s \tag{11}$$

$$B_n^{s,RF'} = B_{n-1}^{s,RF'}\Phi_{ss} - \delta_1^{s'} \tag{12}$$

Risk-adjusted parameters imply that the model-fitted yields equal the time t expectation of the average future short rates over the next n periods, $E_t\left(-\left(\frac{1}{n}\right)lnP_t^{(n)}\right) = -\left(\frac{1}{n}\right)\left(A_n^{s,RF} + B_n^{s,RF'}X_t^s\right)$. The risk neutral yield (RNY), and the term premium (TP), the difference between the model-implied fitted yield and the risk neutral yield, can be written as

$$RNY_t^{(n)} = -\left(\frac{1}{n}\right)\left[A_n^{s,RF} + B_n^{s,RF'}X_t^s\right] \tag{13}$$

$$TP_t^{(n)} = -\left(\frac{1}{n}\right)\left[\left(A_n^s - A_n^{s,RF}\right) + \left(B_n^s - B_n^{s,RF}\right)'X_t^s\right] \tag{14}$$

To investigate how unspanned macroeconomic variables (X_t^u), and the Chinese foreign exchange reserves, in particular, affect the spanned factors (X_t^s) and (log) bond prices ($lnP_{j,t}^{(n)}$) I focus on $\hat{\lambda}_{su}$. I increase principal components extracted from the US yield curve by the estimated coefficients $\hat{\lambda}_{su}$ which are statistically significantly different from zero and correspond to the Chinese foreign reserves. I compare the change in the mean of the model implied 5y Treasury yield, the 5y Treasury risk-neutral yield and the 5y Treasury term premium before and after I increase the principal components by the estimated coefficients $\hat{\lambda}_{su}$. I interpret the difference in the means as the average effect of the Chinese foreign reserves on the model implied 5y Treasury yield, the model implied 5y Treasury risk-neutral yield and the model implied 5y Treasury term premium.

3. Data

I estimate the model at a monthly frequency from December 2000 to December 2017. I focus on the maturities from 1 to 60 months (5 years). The rest of the data is as follows. US CPI inflation and unemployment rate are retrieved from the FRED database of the Federal

Reserve Bank of St. Louis. The US CPI inflation is measured as a log 12-month difference in the seasonally adjusted CPI index. I retrieve Chinese CPI inflation from OECD. Yearly change in Chinese foreign reserves is calculated as a 12-month log difference in Chinese foreign exchange reserves which are presented in Figure 4. The Chinese leading indicator is retrieved from the OECD. The US Dollar index is retrieved from the FRED database.

Figure 2 depicts the 5y Treasury yield in the period from December 2000 to December 2017. The 5y Treasury yield decreased from 5 percent in December 2000 to 2 percent in January 2009. It increased to 2.6 percent in March 2010 most probably due to ongoing Sovereign debt crisis in the Euro Area. In August 2012, it reached a minimum of 0.6 percent. Following the FED's hiking cycle which started in December 2015, the 5y Treasury yield increased to 2.2 percent in December 2017.

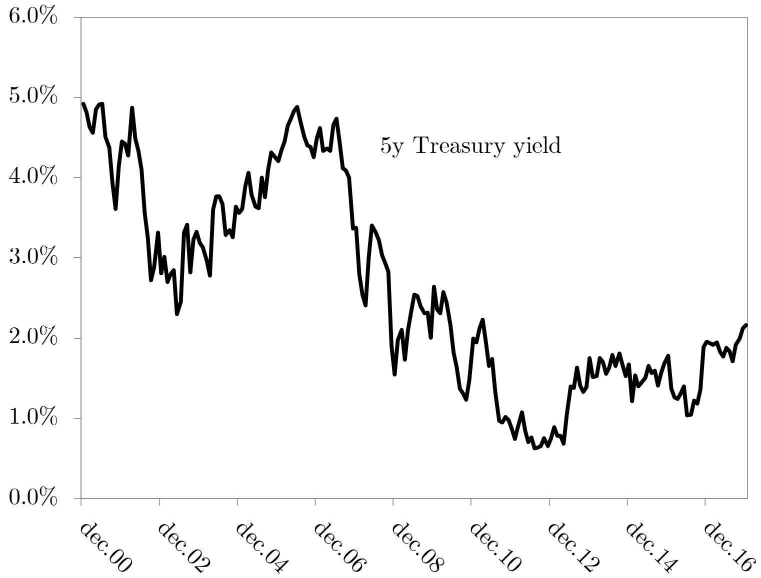


Figure 2: 5y Treasury yield. Sample spans from December 2000 to December 2017.

Figure 3 shows the loadings of the first four principal components extracted from the US term structure on yields of different maturities. Loadings of the first principal component are fairly similar across maturities. Hence, the first principal component is usually denoted as the level factor. The second principal component loads negatively on the yields of shorter maturities and positively as we move to the maturities further out on the yield curve. Therefore, the second principal component is denoted as the slope factor. The loadings of the third principal component resemble the usual shape of curvature. Since the yields used to estimate the affine term structure model are smoothed, the loadings of the fourth principal component on the yields of different maturities are smoothed as well but lack the economic content.

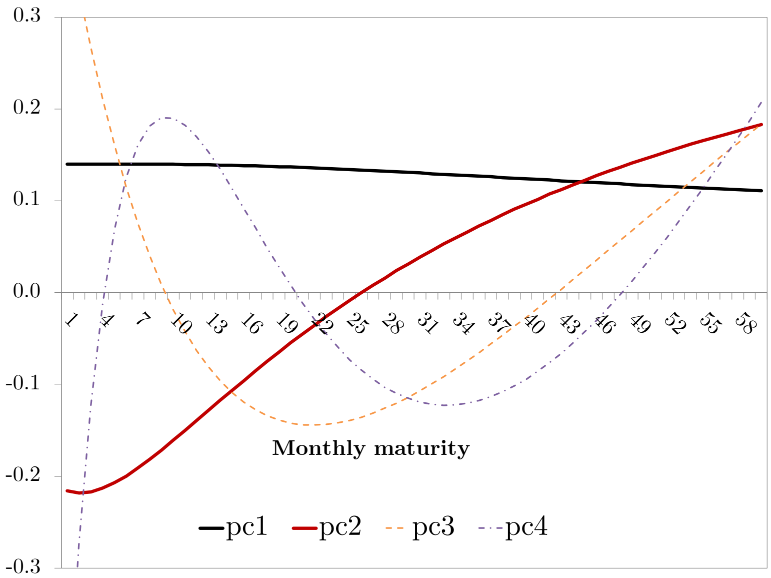


Figure 3: Loadings of principal component 1 (PC1), principal component 2 (PC2), principal component 3 (PC3), and principal component 4 (PC4) on yields of different maturities (60 monthly maturities). Sample spans from December 2000 to December 2017.

In 1998, global allocated foreign reserves equaled 1.5 trillion US Dollars from where they increased to 10 trillion US dollars in 2017. At the end of 2017, more than 60 percent of the total allocated reserves were denominated in US dollars. The relative importance of Chinese foreign exchange reserves in the global reserves increased substantially over time. Figure 4 shows the Chinese foreign exchange reserves. In December 1999, the Chinese foreign reserves equaled 160 million US Dollars. They increased to 3.2 trillion US Dollars by December 2011 and increased further to 4 trillion US Dollars by June 2014. China is the biggest holder of the foreign exchange reserves in the world. In May 2018, the People's Bank of China held 3.1 trillion US dollars of foreign reserves, which is approximately a quarter of total global foreign reserves[17].

[17] Data can be obtained from IMF's COFER database.

Between 1995 and 2010 China acquired roughly $1.1 trillion of US Treasury notes and bonds[18]. The accumulation reflects the Chinese current account surplus and the foreign exchange rate policy. China pegged the Renminbi to the US dollar in 1994. In July 2005, the Chinese authorities implemented a managed floating rate. The daily central parity rate was announced around which the Renminbi would fluctuate. In 2015, the PBOC announced a move towards a market based determination of the Renminbi exchange rate. On the day of the announcement, 11th of August 2015, the Renminbi depreciated by 1.9 percent, and continued to depreciate (by additional 1 percent on 12th of August 2015)[19]. The change of the exchange rate regime was a signal to a market of a prolonged period of the Renminbi depreciation. To prevent the negative effects of the significant Renminbi depreciation the PBOC intervened in the foreign exchange markets and appreciated the Renminbi which has depleted the Chinese foreign exchange reserves.

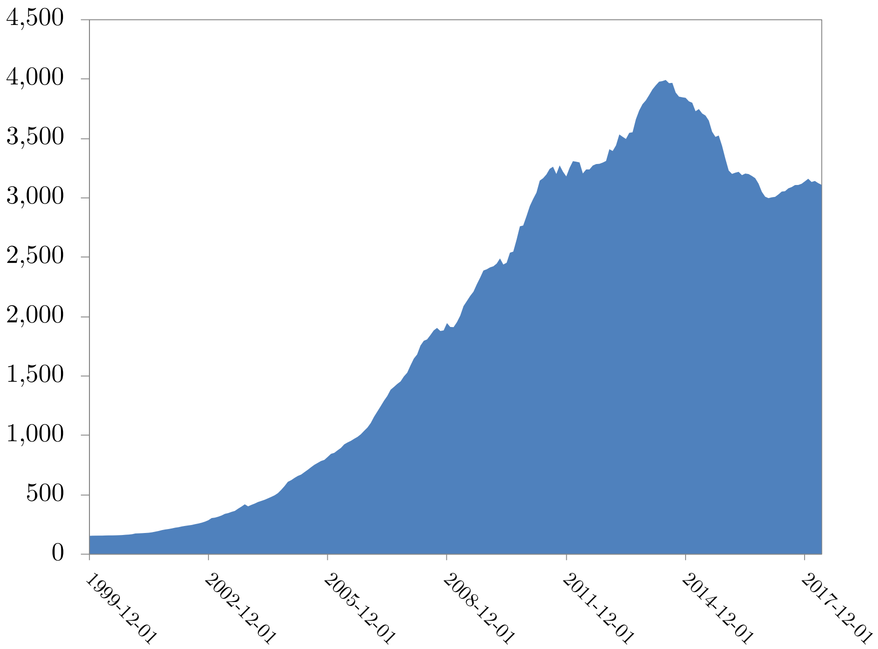


Figure 4: Total foreign exchange reserves of the People's Bank of China (PBOC) in millions of US dollars. Source: Chinese State Administration of Foreign Exchange (SAFE).

[18] Data can be obtained at https://www.treasury.gov/resource-center/data-chart-center/tic/Pages/ticsec.aspx

[19] The announcement is available at http://www.pbc.gov.cn/english/130721/2941603/index.html

Figure 5 depicts the Renminbi against the US Dollar, the *level* of the Dollar index, and the Renminbi nominal effective exchange rate. The Dollar index is the foreign exchange value of the US dollar which is weighted against the major US trading partners[20,21]. We can observe that until 2005 the Renminbi equaled 8.3 US Dollars. In July 2005, the PBOC moved away from the fixed and moved to a managed floating exchange rate regime. Until July 2015, the nominal effective exchange rate of the Renminbi appreciated by 44 percent. In August 2015, the PBOC announced a move towards increased flexibility in setting up the exchange rate which was followed by a managed depreciation. From July 2015 until December 2017 the nominal effective exchange rate of the Renminbi depreciated by 8 percent[22].

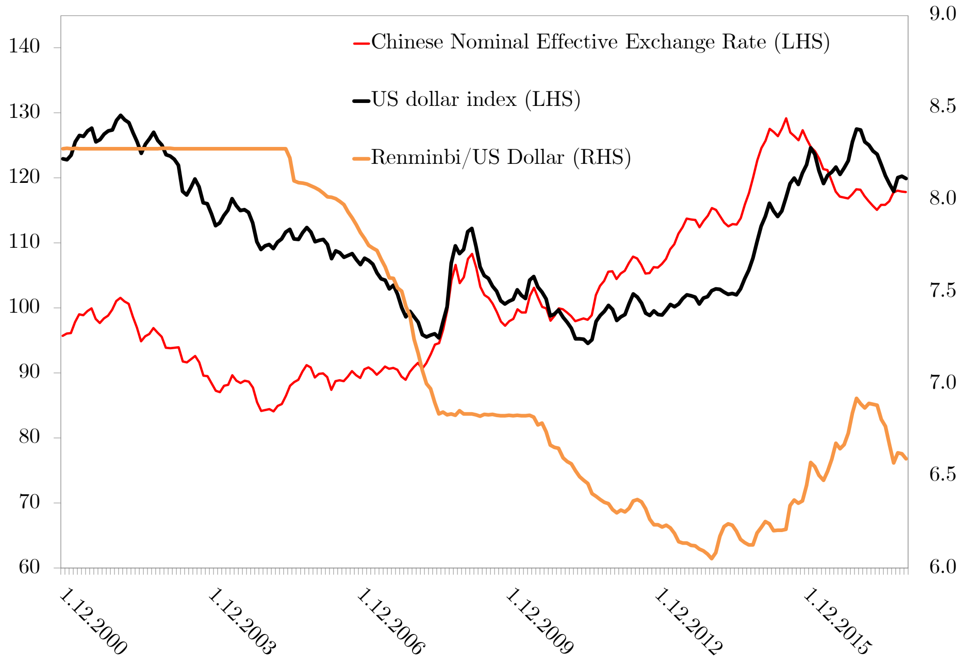


Figure 5: Renminbi/US Dollar, the Dollar index (in levels), and the Chinese nominal effective exchange rate. Series are presented in levels. Sample spans from December 2000 to December 2017. Sources: St. Louis FRED and IMF.

---

[20] The share of China in the index increased from 6.6 percent in 1997 to 16.2 percent in 2017. The correlation between the US Dollar index and the USD/EUR exchange rate in the period from December 1999 to December 2017 equals −0.92. The European Union is the biggest trading partner and represents 18.6 percent of the index in 2017. The index weights are available at https://www.federalreserve.gov/releases/h10/Weights/

[21] Data can be retrieved at https://fred.stlouisfed.org/series/TWEXBMTH

[22] The Chinese nominal effective exchange rate is available at http://data.imf.org/regular.aspx?key=61545850

Figure 6 depicts the Chinese foreign reserves and the Dollar index *in 12-month differences* to provide an intuition of how important is the historical link of the Renminbi to the US Dollar for the development of the Chinese foreign exchange reserves.

The yearly growth of Chinese foreign exchange reserves increased from 7 percent in December 2000 to 53 percent by November 2003. With some minor interruptions, the yearly growth of the Chinese foreign reserves stayed high until the financial crisis. The average yearly growth of Chinese foreign exchange reserves from November 2003 to December 2007 equaled 40 percent. However, during the financial crisis, the growth of foreign reserves decreased substantially. By June 2009, the yearly growth of foreign reserves decreased to 18 percent. The growth increased back to 30 percent by June 2011. It equaled 0 percent in July 2012.

The correlation between the Chinese foreign reserves and the yearly growth rate of the US Dollar index is high in absolute terms, −0.66. The weaker US Dollar pulls down the Renminbi which was pegged to the US Dollar before 2005 and afterwards linked to the US Dollar through the daily central parity rate. When the US Dollar depreciates, among other measures, the PBOC buys the foreign assets and sells the Renminbi in order to depreciate the Renminbi. We see several such episodes of the depreciating US Dollar before the financial crisis.

From December 2000 to November 2003 the growth of the Chinese foreign reserves increased from 7 percent to 53 percent. The yearly growth rate of the US Dollar decreased from 6.3 percent to −8 percent. The US Dollar depreciated by 14.3 percent. From January 2006 to November 2007 the growth of the Chinese foreign reserves increased from 35.5 percent to 44.1 percent. The yearly growth rate of the US Dollar index decreased from 0.6 percent to −8.6 percent. The US Dollar depreciated by 9.2 percent.

However, after the financial crisis, the US Dollar tended to appreciate. From June 2006 to June 2012 the growth of Chinese foreign exchange reserves decreased from 30 percent to 1 percent. The yearly growth rate of the US Dollar index increased from −9.6 percent to 7 percent. The US Dollar appreciated by 16.6 percent. From July 2014 to August 2015 the growth of foreign reserves decreased from 11.8 percent to −10.4 percent. The yearly growth rate of the US Dollar index increased from 0 percent to 14.6 percent. The US Dollar has appreciated. In Figure 4 we can observe that the foreign reserves decreased from 4 trillion US Dollars in June 2014 to 3.1 trillion US Dollars by December 2017. In the same period, the US dollar appreciated by 16 percent.

At the end of my sample, the US Dollar has depreciated. From May 2016 to December 2017 the yearly growth rate of the US Dollar index decreased from 5.4 percent to −6.1 percent. The growth of Chinese foreign reserves increased from −14 percent to 4.3 percent. In Figure 6, we can see that the historical buildup of Chinese foreign reserves is not immune to the development of the US Dollar. The Renminbi was in one way or another linked to the US dollar throughout most of my sample. Correspondingly, when the US Dollar depreciates, the Chinese foreign reserves seem to increase. Especially after the financial crisis, we are witnessing several longer periods of the US Dollar appreciation. Correspondingly, the Chinese foreign exchange reserves have depleted.

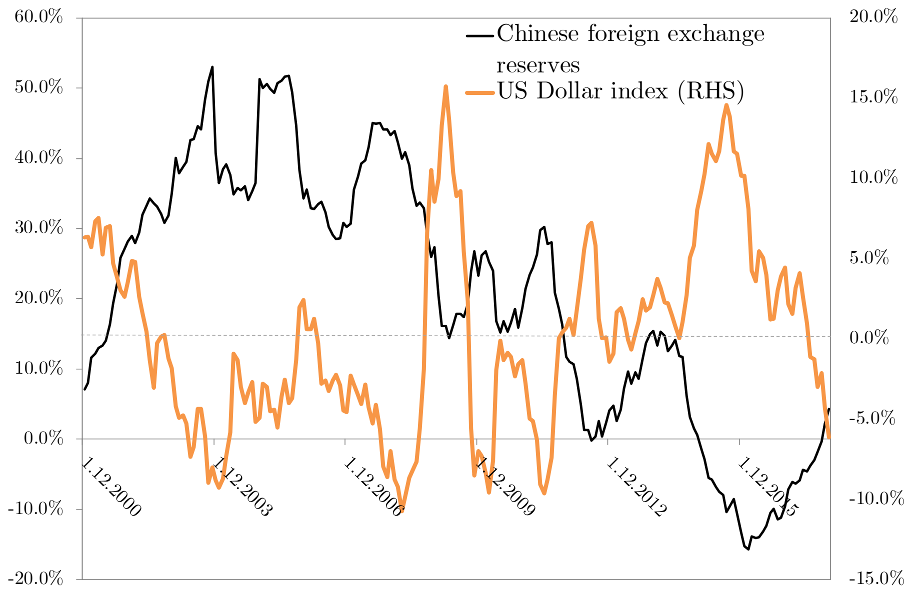


Figure 6: Chinese foreign reserves and the Dollar index. Both series are presented in log 12-month differences. Sample spans from December 2000 to December 2017. Source: St. Louis FRED and the Chinese State Administration of Foreign Exchange (SAFE).

The main contribution of my paper is to estimate the effects of the Chinese foreign reserves on the 5y Treasury yield and the US yield curve. I decompose the 5y Treasury yield in the future expected short rate and the 5y Treasury term premium. Figure 7 depicts the growth of foreign reserves and the 5y Treasury term premium. We can see that both series are in a downward trend. The growth of Chinese foreign reserves increased from 7 percent in December 2000 to 53 percent in November 2003. Afterwards, the growth of the Chinese foreign reserves decreased to 4 percent by December 2017. Over the same period, the 5y Treasury term premium decreased from 1.8 percent in 2003[23] to 0.3 percent in 2017.

The two series have a strong correlation in my sample. The correlation equals 0.69. After the financial crisis, we witnessed at least two episodes when the 5y Treasury term premium decreased at almost the same time as the Chinese foreign reserves. In March 2011, the 5y Treasury term premium locally peaked at 1.2 percent. There months later, the growth of the Chinese foreign reserves peaked at 30.3 percent. Both series reached a local minimum in July 2012. The 5y Treasury term premium decreased to -0.7 percent, and the growth of the Chinese foreign reserves decreased to -1 percent. In December 2013, the 5y Treasury term premium was equal to 0.7 percent. The growth of the Chinese foreign reserves equaled 15.4 percent. Both series decreased by 2016. In January 2016, the 5y Treasury term premium decreased to -0.2 percent. The growth of Chinese foreign reserves decreased to -15.3 percent.

[23] The average 5y Treasury term premium throughout 2003.

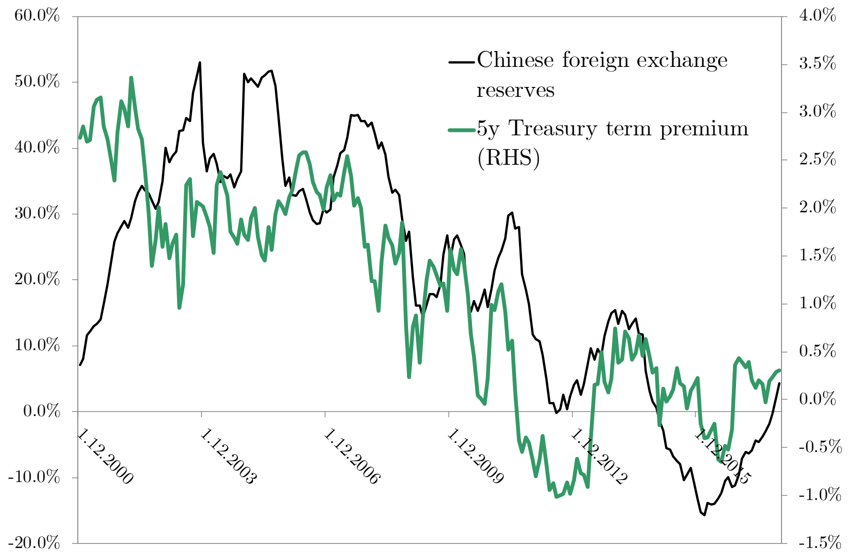


Figure 7: Chinese foreign reserves and the 5y Treasury term premium. Sample spans from December 2000 to December 2017. Source: Chinese State Administration of Foreign Exchange (SAFE) and own calculations. The 5y Treasury term premium is extracted with the four-factor affine term structure model with the unspanned macroeconomic variables as presented in the second section of this paper.

To motivate why it is important to control for the Chinese CPI inflation when measuring the effects of the Chinese foreign exchange reserves on the US yield curve, Figure 8 depicts Chinese inflation and the Chinese foreign exchange reserves from December 2000 to December 2017. The correlation between the series is equal to 0.21. Before the financial crisis, the correlation between the series equaled 0.35. It increased to 0.98 during the financial crisis (from December 2007 to June 2009) and decreased back to 0.43 after the crisis. We can observe that the Chinese CPI inflation increased from 0 percent in 2000 to 8.5 percent in April 2008. During the financial crisis, Chinese inflation decreased substantially. In July 2009, Chinese inflation equaled -1.8 percent. It increased to 6.4 percent in June 2011. Chinese inflation decreased to 1.7 percent in October 2012. After 2012, Chinese inflation became much less volatile. In the period from October 2012 to December 2017, its standard deviation equaled 0.6 percent.

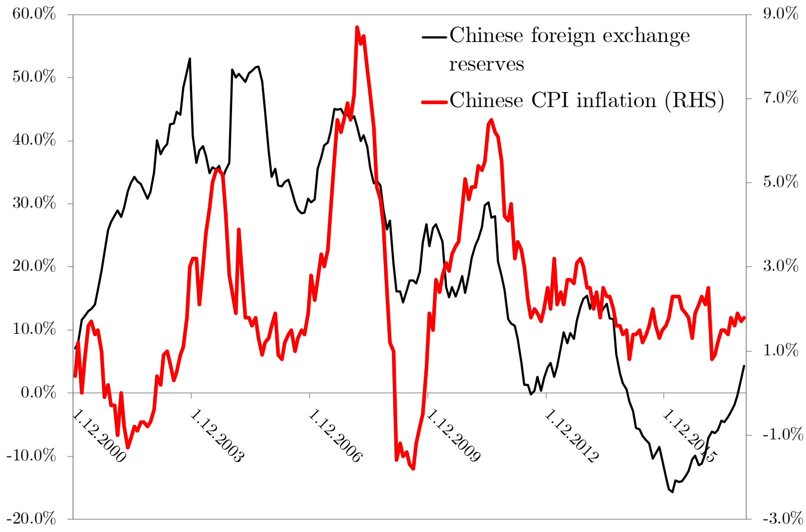


Figure 8: Chinese inflation (CPI index), and the Chinese foreign reserves in log 12-month differences. Sample spans from December 2000 to December 2017. Source: OECD and Chinese State Administration of Foreign Exchange (SAFE).

Although not first order important for the results of this paper, I present the cross-correlations between the Chinese CPI inflation, US CPI inflation, and the rest of the world. In the second column of Table 1, we can see that the correlation between the US and CPI inflation of advanced economies is high and equals 0.91. The correlation between the US and Chinese inflation is much lower, 0.37. However, the monthly correlation between the US and Chinese CPI inflation increases to 0.52 (not presented in Table 1, the series are plotted in Figure 9). The correlation between the Chinese CPI inflation and the inflation of advanced economies is in between these two numbers and equals 0.46. The correlation between Chinese CPI inflation and the CPI inflation of emerging market economies is the lowest, 0.09. Overall, the relatively low correlations between the Chinese CPI inflation and the rest of the world suggest that factors such as the pegged exchange rate and the Chinese foreign exchange rate policy lowered the correlation of Chinese inflation with the rest of the world.

Table 1: Correlations between the Chinese CPI inflation rate, the US CPI inflation rate, the CPI inflation rate of advanced economies, the CPI inflation rate of Emerging market economies and the global CPI inflation. The data are extracted from the IMF's WEO database. Sample spans from 2000 to 2017 (yearly data).

Correlation matrix	CPI_{ch}	CPI_{us}	CPI_{AE}	CPI_{EM}	CPI_{global}
CPI_{ch}	1.000				
CPI_{us}	0.369	1.000			
CPI_{AE}	0.455	0.908	1.000		
CPI_{EM}	0.087	0.368	0.578	1.000	
CPI_{global}	0.386	0.604	0.831	0.891	1.000

Figure 9 depicts the Chinese CPI inflation and the US CPI inflation at a monthly frequency. The US inflation increased from 3.4 percent in December 2000 to 4 percent by December 2007. In the same period, the Chinese CPI inflation increased from 0 percent to 6.5 percent. During the financial crisis, we observed a sharp movement in both series. The swing in Chinese inflation was higher than in the US. At the end of the financial crisis, in June 2009, Chinese inflation decreased to -1.7 percent. US inflation decreased to -1.2 percent. After the crisis, Chinese inflation increased substantially. It was equal to 6.4 percent in June 2011. US inflation increased to 3.4 percent. The standard deviation of US inflation in the period from January 2012 to December 2017 equaled 0.8 percent. The standard deviation of Chinese inflation in the same period was lower and equaled 0.7 percent.

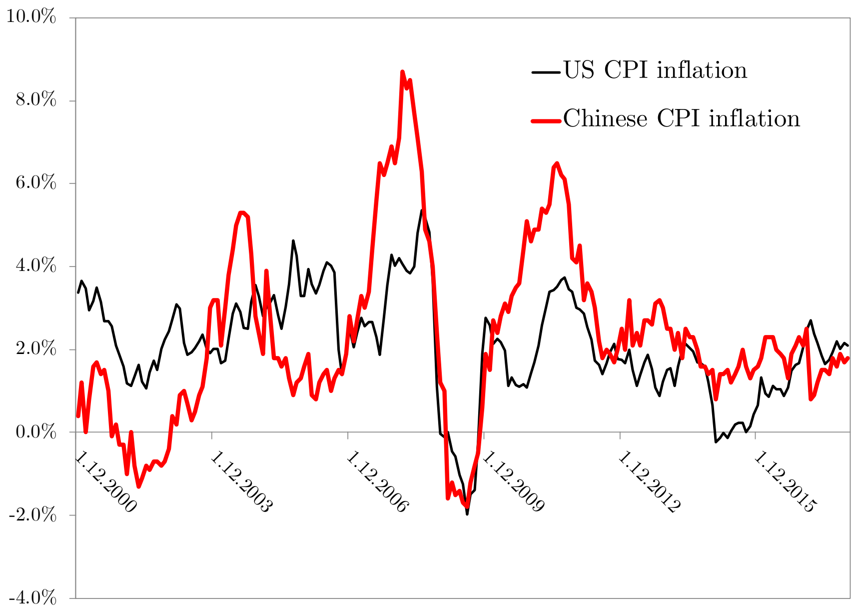


Figure 9: Chinese inflation (CPI index), and the US CPI inflation (log 12-month differences of seasonally adjusted CPI index). Sample spans from December 2000 to December 2017. Source: OECD and St. Louis FRED.

4. Main Results

In this section, I present my main empirical results. Before I introduce the decomposition of the 5y Treasury yield in the 5y Treasury risk-neutral yield and the 5y Treasury term premium, I present the short and long-run effects of the Chinese foreign exchange reserves on the actual 5y Treasury yield and vice-versa. Table 2 presents the estimated short-run effects. In the first row, we can see that the US unemployment and change of Chinese foreign reserves affect the 5y Treasury yield significantly. Albeit, the Chinese foreign reserves merely at the 6.3 percent level. One percentage point higher US unemployment rate decreases the 5y Treasury yield by 4 basis points.

The estimated coefficients in the fourth column of Table 2 represent an increase of the Chinese reserves by “a unit”. In my case, this implies that the Chinese foreign reserves would increase by 100 percent. Therefore, a 10 percent increase of the Chinese foreign reserves increases the 5y Treasury yield by 3.6 basis points[24]. In the fourth row, we see the feedback from the 5y Treasury yield to the Chinese foreign reserves. One percentage point increase of

---

[24] The in-sample mean of the Chinese foreign reserves equals 20.4 percent, and standard deviation of the Chinese foreign reserves equals 18.4 percent.

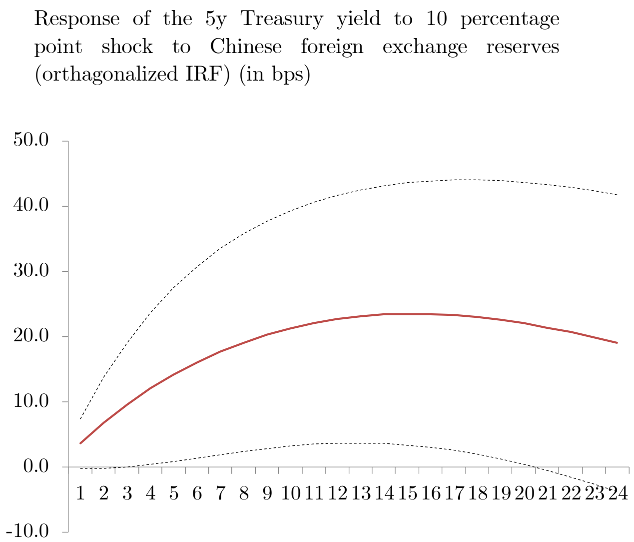
the 5y Treasury yield *increases* the Chinese foreign reserves by 0.81 *percent*. One percentage point increase in the Dollar index *decreases* the Chinese foreign reserves by 0.12 *percent*. When I replace the 5y Treasury yield with the four principal components extracted from the US yield curve, I find that only the estimated coefficient of the first principal component is affecting significantly the Chinese foreign exchange reserves.

Table 2: Estimated coefficients of a five-variable vector autoregression: $X_t = \mu + \Phi X_{t-1} + \varepsilon_t$. Variables included in the regression: 5y Treasury yield, US unemployment, US CPI inflation, Chinese foreign reserves, Chinese CPI inflation, the Dollar index, and the Chinese leading indicator. Sample spans from December 2000 to December 2017. Bolded coefficients are significant at the 10% level.

Factor	$\boldsymbol{\Phi_{1,1}}$	$\boldsymbol{\Phi_{1,2}}$	$\boldsymbol{\Phi_{1,3}}$	$\boldsymbol{\Phi_{1,4}}$	$\boldsymbol{\Phi_{1,5}}$	$\boldsymbol{\Phi_{1,6}}$	$\boldsymbol{\Phi_{1,7}}$
	(5y Treasury Yield)	(ur_{us})	(CPI_{us})	(Foreign reserves)	(CPI_{ch})	(Dollar index)	(CLI_{ch})
5y Treasury Yield	**0.9051**	**-0.0399**	-0.0180	**0.0036**	-0.0169	-0.0018	0.0034
(t-statistic)	29.40	-2.51	-0.79	1.86	-1.40	-0.29	0.45
ur_{us}	**0.0468**	**1.0211**	-0.0154	**0.0035**	0.0006	0.0037	**-0.0174**
(t-statistic)	2.54	107.04	-1.12	3.00	0.08	0.99	-3.80
CPI_{us}	0.0454	**-0.0438**	**0.7988**	**-0.0065**	**0.0456**	**-0.0307**	**0.0241**
(t-statistic)	0.90	-1.68	21.39	-2.04	2.32	-3.04	1.92
Foreign reserves	**0.8055**	0.0338	-0.3444	**0.9330**	-0.0793	**-0.1195**	0.0513
(t-statistic)	2.70	0.22	-1.56	49.46	-0.68	-1.99	0.69
CPI_{ch}	-0.0124	-0.0516	**-0.0860**	**-0.0099**	**0.9571**	**-0.0318**	**0.0524**
(t-statistic)	-0.18	-1.46	-1.69	-2.30	35.73	-2.31	3.08
Dollar index	-0.0386	**0.2011**	**0.4395**	0.0204	-0.1140	**0.9382**	**-0.1982**
(t-statistic)	-0.19	1.92	2.92	1.59	-1.43	23.02	-3.93
CLI_{ch}	**0.1278**	**0.0748**	**-0.1750**	**0.0121**	**-0.1072**	**0.0324**	**0.9735**
(t-statistic)	2.67	3.02	-4.92	3.98	-5.71	3.37	81.65

Figure 10 depicts the long run responses of the 5y Treasury yield to a unit shock to Chinese foreign reserves (left panel). The response of the 5y Treasury yield to a 10 percentage point increase of the Chinese foreign exchange reserves equals 3.6 basis points in the 1st month. I impose the Cholesky decomposition with the lower triangular variance-covariance matrix of shocks and order the Chinese foreign reserves last. While other variables can contemporaneously affect the Chinese foreign reserves, the foreign reserves are allowed only to react to other variables included in the vector-autoregression with a lag. In the 14th month, the response of the 5y Treasury yield peaks and increases to 23.5 basis points.

The right panel of Figure 10 shows the responses of the Chinese foreign reserves to a one percentage point shock to the 5y Treasury yield (the latter is ordered last). The foreign reserves increase by 0.80 percentage point in the 1st month. The effect increases to 4.76 *percentage points* in the 23th month. In the 24th month, the response decreases to 4.75 percentage points and is still significantly different from 0. I perform a similar analysis to calculate the long-run responses of the US Dollar index. The responses become insignificant after the 1st month.



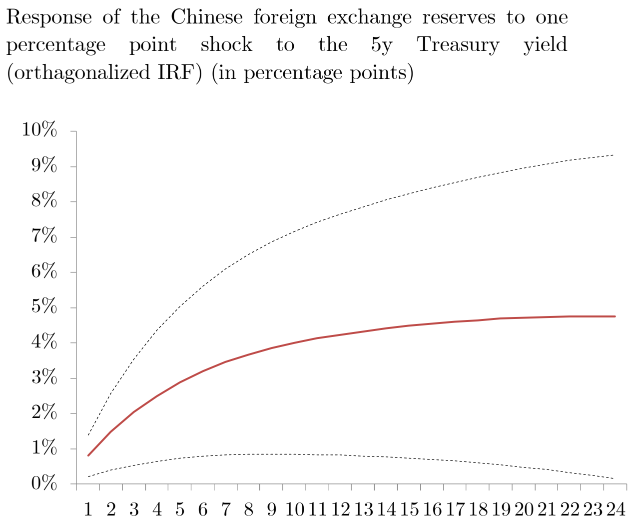


Figure 10: Orthogonalized impulse response functions of the 5y Treasury yield to 10 percentage point shock to Chinese foreign exchange reserves (left panel) in a structural vector auto-regression with Cholesky identification scheme. The right panel shows the responses of the Chinese foreign exchange reserves to 1 percentage point shock to the 5y Treasury yield. Sample spans from December 2000 to December 2017. Variables included in the model: 5y Treasury yield, US unemployment, US CPI inflation, Chinese foreign exchange reserves, Chinese CPI inflation, Chinese leading indicator, and the Dollar index. In the left panel, Chinese foreign exchange reserves are ordered last, and in the right panel, the 5y Treasury yield is ordered last. The lower-triangular variance-covariance matrix of shocks is imposed.

Decomposition of the 5y Treasury yield in the 5y Treasury risk-neutral yield and the 5y Treasury term premium

I estimate the affine term structure model of the US yield curve with the unspanned macro variables which are presented in Section 2 in the period from December 2000 to December 2017 following Diez de Los Rios (2018). Table 3 presents the estimated prices of risks of the four-factor model. The level risk is priced by itself, the second, third principal component, and the Chinese foreign exchange reserves. The slope risk is priced only by the second principal component and the Chinese foreign reserves. The third principal component is priced by the first, the second principal component, and the Chinese foreign reserves. The risk of the fourth principal component is priced by itself, the second principal component, and the Chinese foreign reserves. Therefore, Chinese foreign reserves affect all four principal components significantly. Among four different specifications, one-, two-, three-, and four-factors, the four-factor model has the lowest average pricing error of the 5y Treasury yield, 0.8 basis points.

In Figure 10 we can see that feedback from the 5y Treasury yield to the foreign reserves is positive. Higher 5y Treasury yield implies lower bond prices. Since US Treasuries are an important part of the Chinese foreign exchange reserves, higher 5y Treasury yield should lower and not increase the growth of the Chinese foreign exchange reserves. I find that only the first principal component extracted from the US yield curve affects the growth of the Chinese foreign exchange reserves significantly.

If the (uncovered) interest rate parity holds higher US nominal interest rates depreciate the US Dollar in the future (assuming that the real exchange rate remains unchanged). This increases Chinese foreign reserves. The economic mechanism helps to explain a relatively important contribution of the level factor for movements in the 5y Treasury term premium. A unit increase of the level factor increases the 5y Treasury yield by 13.2 basis points, the 5y Treasury term premium by 11 basis points and the 5y Treasury risk-neutral yield by 2.3 basis points.

Table 3: Estimated prices of risk, λ_0^s and $\lambda_1^{s\cdot}$ in the affine term structure model as outlined in Diez de Los Rios (2018). Sample spans from December 2000 to December 2017. Spanned factors: $X_t^s = [PC\ 1_t \quad PC\ 2_t \quad PC\ 3_t \quad PC\ 4_t]'$. Unspanned factors: $X_t^u = [ur_{US,t} \quad CPI_{US,t} \quad FXR_{CH,t} \quad CPI_{CH,t} \quad Dollar\ Index_t \quad CLI_{CH,t}]'$. $ur_{US,t}$ – US unemployment rate, $CPI_{US,t}$ – US CPI inflation rate, $FXR_{CH,t}$ – Chinese foreign exchange reserves, $CPI_{CH,t}$ – Chinese CPI inflation rate, $Dollar\ Index_t$ – The Dollar index, and $CLI_{CH,t}$ – Chinese leading indicator. Bolded coefficients are significant at the 5% level. I present the remaining estimated parameters in the Appendix A.1..

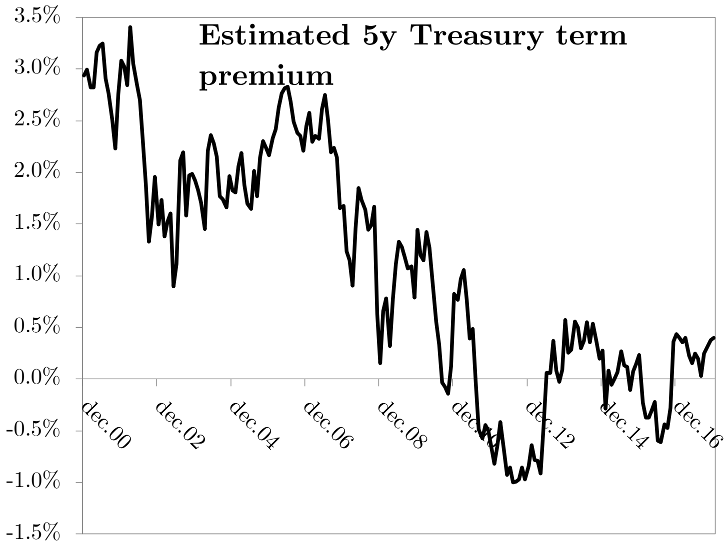
Factor	$\boldsymbol{\lambda_0}$	$\boldsymbol{\lambda_{1,1}}$	$\boldsymbol{\lambda_{1,2}}$	$\boldsymbol{\lambda_{1,3}}$	$\boldsymbol{\lambda_{1,4}}$	$\boldsymbol{\lambda_{1,5}}$	$\boldsymbol{\lambda_{1,6}}$	$\boldsymbol{\lambda_{1,7}}$	$\boldsymbol{\lambda_{1,8}}$	$\boldsymbol{\lambda_{1,9}}$	$\boldsymbol{\lambda_{1,10}}$
	(constant)	(PC1)	(PC2)	(PC3)	(PC4)	(ur_{us})	(CPI_{us})	(FXR_{ch})	(CPI_{ch})	(Dollar index)	(CLI_{ch})
PC 1	**0.3161**	**-0.0544**	**-0.2873**	**-0.7532**	-0.8161	-1.3202	-1.0467	**0.2432**	-1.7183	-0.3543	0.0036
(t-statistic)	4.41	-2.39	-4.11	-3.06	-0.87	-1.31	-1.60	2.30	-1.36	-1.08	0.01
PC 2	**0.0707**	-0.0023	**-0.1248**	0.0290	0.0907	-0.1659	-0.3770	**0.1065**	-0.8660	-0.0341	-0.2338
(t-statistic)	2.26	-0.23	-4.08	0.27	0.22	-0.38	-1.32	2.31	-1.57	-0.24	-1.25
PC 3	-0.0123	**0.0081**	**0.0433**	-0.0651	-0.1142	0.2523	-0.0908	**-0.0400**	0.1054	-0.0848	-0.0500
(t-statistic)	-1.02	2.12	3.69	-1.57	-0.73	1.50	-0.83	-2.26	0.50	-1.55	-0.70
PC 4	0.0044	-0.0009	**-0.0197**	-0.0212	**-0.2162**	0.0954	0.0583	**0.0165**	-0.1116	0.0007	0.0057
(t-statistic)	0.99	-0.64	-4.51	-1.37	-3.69	1.55	1.46	2.55	-1.44	0.04	0.22

Figure 11 depicts the estimated 5y Treasury term premium (upper panel) and the estimated 5y Treasury Risk-Neutral yield (lower panel) which are estimated with the unspanned affine term structure model. In December 2000, the 5y Treasury term premium equaled 2.9 percent. It steadily decreased to 0 percent in August 2010. The 5y Treasury yield, over the same period, decreased from 4.9 percent to 1.4 percent. The 5y Treasury risk-neutral yield decreased from 2.1 percent in December 2000 to 1.4 percent (lower panel of Figure 11).

The 5y Treasury term premium increased from -0.1 percent in October 2010 to 1.1 percent in March 2011. The 5y Treasury term premium decreased to -1 percent in August 2012. This period coincides with the initiation of the QE programmes by the ECB. In December 2013, the 5y Treasury term premium peaked at 0.6 percent but overall stayed extremely low in the period from December 2014 to December 2017. At the end of my sample, in December 2017, the 5y Treasury term premium equaled 0.4 percent.

The dotted red line in the lower panel of Figure 11 depicts the 1m Treasury yield. We can see that the dynamics of the 1m Treasury yield and the 5y Treasury risk-neutral yield are similar. Before the financial crisis, the 5y Treasury risk-neutral yield stayed most of the time below the 1m Treasury yield. The 5y Treasury risk-neutral yield decreased from 2.1 percent in December 2000 to 1.3 percent in March 2003. It increased to 2.1 percent in July 2006. In July 2006, the 1m Treasury yield was equal to 5.1 percent.

After the financial crisis, the 5y Treasury risk-neutral yield is consistently above the 1m Treasury yield. In June 2009, when the NBER officially announced the end of the recession, the 5y Risk-Neutral yield was equal to 1.2 percent. The 1m Treasury yield was equal to 0.1 percent. Between June 2009 and September 2015, the 1m Treasury yield was moving in a narrow band around zero percent. The 5y Treasury risk-neutral yield, on the other hand, increased to 1.6 percent in October 2012 and decreased to 1.2 percent in December 2013. After 2013, the 5y Treasury risk-neutral yield steadily increased. At the end of my sample, in December 2017, the 5y Treasury risk-neutral yield increased to 1.8 percent. The 1m Treasury yield was equal to 1.6 percent and was only 0.2 percent lower than the 5y Treasury risk-neutral yield.



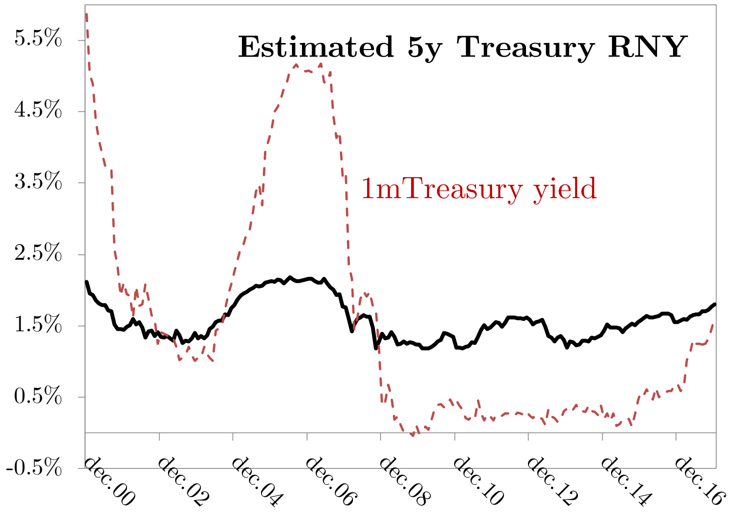


Figure 11: Model implied 5y Treasury term premium and the 5y Treasury expected future nominal short rate (risk neutral yield) estimated with the four-factor model (which uses PC1 to PC4) and unspanned macroeconomic variables: US unemployment, US CPI inflation, Chinese foreign exchange reserves, Chinese CPI inflation, the Dollar index, and the Chinese leading indicator. I use an estimator as outlined in Diez de Los Rios (2018). Sample spans from December 2000 to December 2017.

To measure the effects of the Chinese foreign reserves on the 5y Treasury risk-neutral yield and the 5y Treasury term premium I increase the principal components by the significant estimated coefficients which correspond to the Chinese foreign reserves, $\hat{\lambda}_{1,7}$, in Table 3. The model implied 5y Treasury yield increases by 51 basis points, the implied 5y Treasury term premium by 49 basis points and the implied 5y Treasury risk-neutral yield by 2 basis points. Again, the estimated coefficients represent an increase of the Chinese foreign reserves by "a unit". In my case, this implies that the Chinese foreign reserves would increase by 100 percent. I divide the estimated effects by 10.

A 10 percentage point increase of the Chinese foreign reserves increases the model implied 5y Treasury yield by 5.1 basis points, the model implied 5y Treasury term premium by 4.9 basis points and the model implied 5y Treasury risk-neutral yield by 0.2 basis points. When additionally, I condition on the effective nominal Renminbi exchange rate, the Renminbi against the US Dollar and the VIX, the model implied 5y Treasury yield increases by 6.2 basis points, the model implied 5y Treasury term premium by 5.2 basis points and the model implied 5y Treasury risk-neutral yield by 1 basis points.

The 95 percent confidence interval (in basis points) for 10 percentage points increase of the Chinese foreign exchange reserves of the model implied 5y Treasury yield is $(0.8,\ 9.1)$, the model implied 5y Treasury term premium is $(0.81,\ 8.7)$ and of the model implied 5y Treasury risk-neutral yield is $(-0.01,\ 0.4)$.

When additionally I condition on the Renminbi nominal effective exchange rate, the Renminbi against the US Dollar and the VIX index, the 95 percent confidence interval (in basis points) for 10 percentage points increase of the Chinese foreign exchange reserves of the model implied 5y Treasury yield is $(1.8,\ 10.5)$, the model implied 5y Treasury term premium is $(1.4,\ 9.6)$ and of the model implied 5y Treasury risk-neutral yield is $(0.4,\ 0.9)$.

Economically, I can interpret the estimated effects on the 5y Treasury risk-neutral yield and the 5y Treasury term premium as the impacts on the policy and the risk compensation channels. My empirical findings suggest that the lower growth of the Chinese foreign reserves after the financial crisis seems to be more important for the 5y Treasury yield and the 5y Treasury term premium than the increase of the Chinese foreign reserves before the financial crisis (the 5y Treasury term premium and the growth of the Chinese foreign reserves are depicted in Figure 7). The lower growth of the Chinese foreign reserves signals the lower 5y Treasury yield and decreases the compensation for bearing the duration risk (the 5y Treasury term premium).

Long-run effects on the 5y Treasury risk-neutral yield and the 5y Treasury term premium

To measure the long run effects of the Chinese foreign reserves on the 5y Treasury risk-neutral yield and the 5y Treasury term premium, in the last part of this section, I calculate the responses of the principal component 1 (PC1), the principal component 2 (PC2), the principal component 3 (PC3), and the principal component 4 (PC4) to "a unit" shock to Chinese foreign reserves. I condition on the US unemployment, US CPI inflation, Chinese CPI inflation, the US Dollar, and the Chinese leading indicator which I use in my companion paper (Maletic, 2018) to empirically represent the growth of the Chinese economy. Figure 12 presents the responses.

The PC1 increases by 0.006 in the first month. The response increases to 0.064 in the 24th month. The principal components do not have units and it is, therefore, difficult to argue if these effects are economically important or not. To measure the economic importance of the estimated responses of the PC1, I increase the PC1 by the estimated response in the 24th month, 0.064, and calculate the change in the average model implied 5y Treasury yield, the 5y Treasury term premium, and the 5y Treasury risk-neutral yield.

However, I identify a shock to the Chinese foreign reserves by imposing the Cholesky decomposition of the variance-covariance matrix and order the Chinese foreign reserves last (with the lower triangular variance-covariance matrix). I should, therefore, divide the response by the element in the Cholesky matrix which corresponds to the Chinese foreign reserves, a standard deviation implied by the Cholesky decomposition which is by definition equal to the response of the Chinese foreign reserves on itself in period zero. When I calculate the impulse-response function of the Chinese foreign reserves on itself, the Chinese foreign exchange reserves increase by 0.025 in period zero.

Since I impose the Cholesky decomposition and order the Chinese foreign reserves last, the remaining variables included in the vector-autoregression are not affected in period zero. This can be checked by noticing that the column elements corresponding to the Chinese foreign reserves in the Cholesky matrix are equal to zero for all remaining variables which are included in the model. Hence, to quantify the economic importance of the shock to the

Chinese foreign exchange reserves on the US yield curve, I divide the estimated effects by 0.025 and multiply them by 0.1 (to scale a shock to a 10 percentage point increase of the Chinese foreign reserves).

In a model which includes the US unemployment rate, US CPI inflation, Chinese CPI inflation, US Dollar index, Chinese leading indicator, and Chinese foreign reserves, a 10 percentage point increase of the Chinese foreign reserves increases the model implied 5y Treasury yield by 34.9 basis points, the model implied 5y Treasury term premium by 28.8 basis points and the 5y Treasury risk-neutral yield by 6.1 basis points in the 24th month.

When additionally I condition on the effective nominal Renminbi effective exchange rate, the Renminbi against the US Dollar and VIX, a 10 percentage point increase of the Chinese foreign reserves increases the model implied 5y Treasury yield by 28.2 basis points, the model implied 5y Treasury term premium by 21.9 basis points and the 5y Treasury risk-neutral yield by 6.3 basis points in the 24th month.

The 95 percent confidence interval (in basis points) in the 24$^{\text{th}}$ month for 10 percentage points increase of the Chinese foreign exchange reserves of the model implied 5y Treasury yield is (10.2, 59.6), the model implied 5y Treasury term premium is (8.4, 49.2) and of the model implied 5y Treasury risk-neutral yield is (1.8, 10.4).

When additionally I condition on the Renminbi nominal effective exchange rate, the Renminbi against the US Dollar and the VIX index, in the 24$^{\text{th}}$ month, the 95 percent confidence interval (in basis points) for 10 percentage points increase of the Chinese foreign exchange reserves of the model implied 5y Treasury yield is (3, 53.4), the model implied 5y Treasury term premium is (2.3, 41.4) and of the model implied 5y Treasury risk-neutral yield is (0.7, 12).

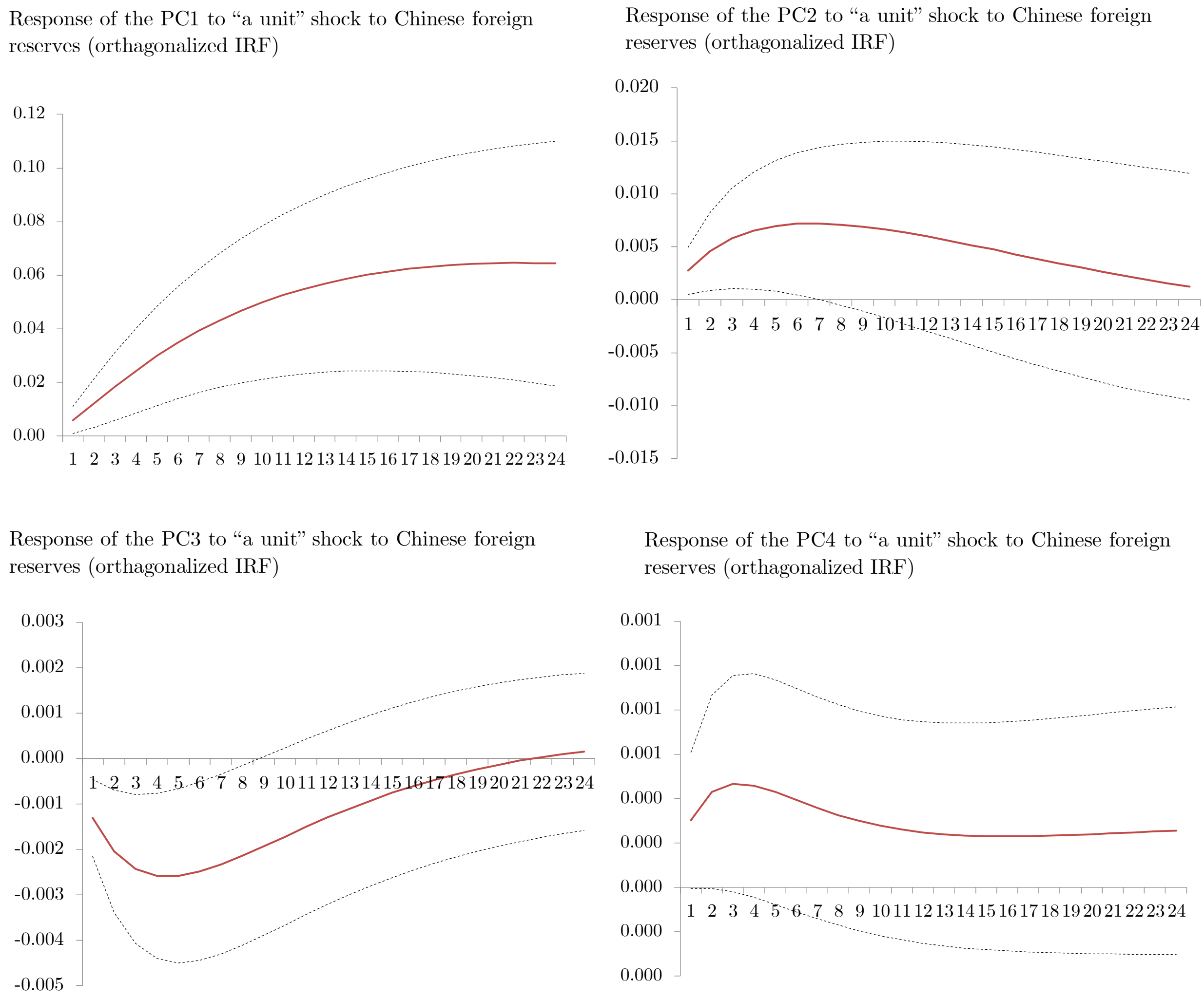


Figure 12: Orthogonalized impulse response functions of the first PC (upper left panel), the second PC (upper right panel), the third PC (lower left panel), and the fourth PC (lower right panel) to "a unit" shock to the Chinese foreign exchange reserves. Sample spans from December 2000 to December 2017. Variables included in the model: PC1 to PC4, US unemployment, US CPI inflation, Chinese foreign exchange reserves, Chinese CPI inflation, US Dollar index and Chinese leading indicator. The Chinese foreign exchange reserves are ordered last. The lower-triangular variance-covariance matrix of shocks is imposed.

5. Conclusions

I quantify the effects of the growth of the Chinese foreign reserves on the 5y Treasury yield, the 5y Treasury risk-neutral yield, and the 5y Treasury term premium. I estimate the affine term structure model with unspanned macroeconomic variables following Diez de Los Rios (2018) in the period from December 2000 to December 2017 which includes the Chinese foreign exchange reserves. The baseline model includes the four principal components extracted from the US yield curve, the US unemployment rate, US CPI inflation, Chinese inflation, the Chinese foreign exchange reserves, the Dollar index, and the Chinese leading indicator.

The lower Chinese foreign reserves represent additional information to the inflation and growth rate of the Chinese economy, and signal a lower 5y Treasury yield and lower the 5y Treasury term premium. China is managing its exchange rate against the US Dollar. My empirical findings suggest that the lower growth of the Chinese foreign reserves after the financial crisis is more important for understanding the low 5y Treasury yield and its term premium than the high growth of the Chinese foreign reserves before the financial crisis.

I find that a 10 percentage point decrease of the Chinese foreign reserves decreases the in-sample average of the model implied 5y Treasury yield by 5.1 basis points, the 5y Treasury term premium by 4.9 basis points and the 5y Treasury risk-neutral yield by 0.2 basis points over the short run. The effects strengthen over the long run. In the 24th month, the 5y Treasury yield decreases by 34.9 basis points, the 5y Treasury term premium by 28.8 basis points, and the 5y Treasury risk-neutral yield by 6.1 basis points. In the 24th month, the 5y Treasury yield decreases by 28.2 basis points, the 5y Treasury term premium by 21.9 basis points, and the 5y Treasury risk-neutral yield by 6.3 basis points, when additionally I condition on the nominal Renminbi effective exchange rate, the exchange rate of the Renminbi against the US Dollar, and the VIX.

When the growth of the foreign reserves of the economy which is as big as China, is an important player in global commodity markets, manages its exchange rate against the US Dollar, and holds a substantial amount of US Treasuries decreases, it signals a lower 5y Treasury yield and decreases the compensation for bearing the duration risk (the 5y Treasury term premium). The risks related to the (unexpected) significant depreciation of the Renminbi against the US Dollar and the lower future growth of the Chinese economy are increasing the risks of lower future global growth and inflation, manifest in the lower Chinese foreign reserves, and decrease the 5y Treasury yield and its term premium.

The economically important feedbacks are running from the US yield curve to the Chinese foreign reserves. The higher level of US nominal interest rates increases the growth of the Chinese foreign reserves. Under the (uncovered) interest rate parity, higher US nominal interest rates depreciate the US Dollar in the future. The PBOC accumulates US Treasuries to depreciate the Renminbi.

References

1. Beltran, D. O., M. Kretchmer, J. Marquez, and C. P. Thomas, 2013, Foreign holdings of US Treasuries and US Treasury yields, *Journal of International Money and Finance*, 32: 1120-1143.
2. Bernanke, B. S., 2015, Why are interest rates so low, part 4: Term premiums, *Brookings blog post.*
3. Bernanke, B. S., 2017, Federal reserve policy in an international context, *IMF Economic Review*, 65(1), 1-32.
4. Bini Smaghi, L., 2010, Reserve Accumulation: The Other Side of the Coin, *Speech at the 5th High-level EMEAP-Eurosystem Seminar.*
5. Das, S., 2019, China's Evolving Exchange Rate Regime, *International Monetary Fund, Working paper No. 19/50.*
6. Department of the Treasury, 2018, Report on Foreign Portfolio Holdings of US Securities as of June 30, 2017, http://ticdata.treasury.gov/Publish/shla2017r.pdf
7. Diez de Los Rios, A., 2015, A new linear estimator for Gaussian dynamic term structure models, *Journal of Business & Economic Statistics*, 33(2), 282-295.
8. Diez de Los Rios, A., 2018, Optimal Estimation of Multi-Country Gaussian Dynamic Term Structure Models Using Linear Regressions, *Working paper.*
9. Frankel, J., 2015, Misinterpreting Chinese Intervention in Financial Markets, *China-U.S. Focus.*
10. Gauvin, L., and Rebillard, C. C., 2015, Towards recoupling? Assessing the global impact of a Chinese hard landing through trade and commodity price channels, *The World Economy.*
11. Kroeber, A., 2011, The Renminbi: The political economy of a currency, *Foreign policy*, 7.
12. Maletic, M., 2018, A Chinese slowdown and the US and German yield curves, *Working paper.*
13. Warnock, F. E., and V. C. Warnock, 2009, International capital flows and US interest rates, *Journal of International Money and Finance*, 28(6), 903-919.

A.1. Estimated parameters of the affine model with the unspanned macroeconomic variables of the US yield curve from December 2000 to December 2017 following Diez de Los Rios (2018)

δ_0^s
(constant)
0.0002
0.0000
(standard error)

Factor	δ_1^s
PC 1	**0.1400**
(standard error)	0.0000
PC 2	**-0.2173**
(standard error)	0.0001
PC 3	**0.4026**
(standard error)	0.0004
PC 4	**-0.4734**
(standard error)	0.0013

Factor	$\mu_s^\star$
	(constant)
PC 1	**0.0003**
(standard error)	0.0000
PC 2	**0.0006**
(standard error)	0.0001
PC 3	**-0.0001**
(standard error)	0.0000
PC 4	**0.0003**
(standard error)	0.0001

Factor	$\Phi_{ss,1,1}^\star$	$\Phi_{ss,1,2}^\star$	$\Phi_{ss,1,3}^\star$	$\Phi_{ss,1,4}^\star$
	(PC1)	(PC2)	(PC3)	(PC4)
PC 1	**0.9941**	**0.1083**	**-0.2222**	**0.3683**
(standard error)	0.0000	0.0001	0.0003	0.0010
PC 2	**-0.0032**	**1.0039**	**0.2672**	**-0.3225**
(standard error)	0.0000	0.0001	0.0004	0.0013
PC 3	**0.0003**	**-0.0265**	**0.8872**	**0.4538**
(standard error)	0.0000	0.0001	0.0004	0.0015
PC 4	**0.0006**	**0.0119**	**0.0196**	**0.8134**
(standard error)	0.0000	0.0001	0.0005	0.0016

	$Eigenvalues$
λ_1	0.8145
λ_2	0.9124
λ_3	0.9671
λ_4	1.0045

Factor	$\Sigma_{1,1}^{Chol}$	$\Sigma_{1,2}^{Chol}$	$\Sigma_{1,3}^{Chol}$	$\Sigma_{1,4}^{Chol}$	$\Sigma_{1,5}^{Chol}$	$\Sigma_{1,6}^{Chol}$	$\Sigma_{1,7}^{Chol}$	$\Sigma_{1,8}^{Chol}$	$\Sigma_{1,9}^{Chol}$	$\Sigma_{1,10}^{Chol}$
	(PC1)	(PC2)	(PC3)	(PC4)	(ur_{us})	(CPI_{us})	(FXR_{ch})	(CPI_{ch})	(Dollar index)	(CLI_{ch})
PC 1	**0.1284**									
(standard error)	0.0067									
PC 2	**0.0379**	**0.0428**								
(standard error)	0.0034	0.0028								
PC 3	**-0.0094**	-0.0009	**0.0197**							
(standard error)	0.0015	0.0018	0.0011							
PC 4	**0.0018**	-0.0001	**-0.0019**	**0.0075**						
(standard error)	0.0006	0.0006	0.0006	0.0004						
ur_{us}	-0.0002	0.0001	-0.0001	0.0001	**0.0014**					
(standard error)	0.0001	0.0002	0.0001	0.0001	0.0001					
CPI_{us}	-0.0001	0.0006	0.0003	-0.0003	-0.0006	**0.0055**				
(standard error)	0.0004	0.0004	0.0004	0.0004	0.0006	0.0002				
FXR_{ch}	-0.0012	0.0012	-0.0005	0.0011	-0.0001	-0.0008	**0.0242**			
(standard error)	0.0018	0.0020	0.0019	0.0019	0.0025	0.0018	0.0012			
CPI_{ch}	**0.0005**	0.0007	0.0003	0.0004	-0.0002	**0.0004**	0.0002	**0.0039**		
(standard error)	0.0005	0.0005	0.0005	0.0005	0.0008	0.0004	0.0005	0.0005		
Dollar index	0.0012	-0.0022	0.0015	0.0012	-0.0011	0.0002	**-0.0074**	**-0.0048**	**0.0136**	
(standard error)	0.0010	0.0011	0.0011	0.0011	0.0018	0.0009	0.0010	0.0017	0.0005	
CLI_{ch}	0.0001	0.0001	-0.0003	0.0001	0.0003	-0.0003	**0.0007**	0.0001	**-0.0006**	**0.0035**
(standard error)	0.0003	0.0003	0.0003	0.0003	0.0003	0.0002	0.0003	0.0004	0.0002	0.0002

Chapter 3: R&D Investments, Past Returns, and the Cross-Section of Stock Returns

Matjaz Maletic[1]

This version: 26th of September 2019

Abstract

This paper investigates how R&D investments and past returns interact in explaining future returns. I gather empirical evidence showing that firms with a higher level of R&D investments, which have experienced lower past returns, are rewarded with higher future returns. Interaction with past returns has the opposite effect on future returns when a firm changes its R&D investments. Firms which increase their R&D investments are rewarded with higher future returns only if they have earned higher past 1-year returns. I am, however, unable to provide empirical support that firms with higher R&D investments, which have experienced lower past returns, are explaining higher future returns of firms with higher R&D-to-market value ratios.

1. Introduction

Chan et al. (2001) provide empirical evidence showing that firms with high R&D-to-market value ratios are rewarded with higher future returns. The R&D-to-market value ratio (RDM) is defined as:

$$RDM_{i,t} = \frac{R\&D\ Expenditures_{i,t}}{Market\ Value_{i,t}} \tag{1}$$

R&D-to-market value ratio, however, is not measuring only the predictive power of R&D investments for the future stock returns. In its denominator, the market value of a company includes share price which affects past return. Chan et al. (2001) find that firms with a high level of R&D investments earn higher future returns only if they have experienced low past returns[2].

---

[1] Tilburg University, Finance Department, and Bank of Slovenia, Research Department. E-mail: matjaz.maletic@bsi.si. I would like to thank Joost Driessen, Frank de Jong and Peter de Goeij for their valuable inputs during my PhD pre-defense which considerably improved the quality of this paper.

[2] Literature which has investigated the relationship between R&D investments and the future stock returns is amongst others (but surely not limited to): Eberhart et al. (2004), Chambers et al. (2002), Aboody and Lev (2000), Lev and Sougiannis (1996, 1999), and Sougiannis (1994).

In this paper, I gather the empirical evidence showing that level and changes in R&D investments interact differently with past returns when quantifying the effect on future stock returns. Firms which have a high level of R&D investments (high R&D-to-assets ratio) are rewarded with higher future returns if they have experienced lower past returns. A firm which changes (increases) R&D investments, on the other hand, is rewarded with higher future returns only if it has experienced higher past 1-year returns. I establish my key results by estimating Fama-Macbeth cross-sectional regressions and testing the predictive power of the interaction term of R&D-to-assets ratio and the yearly growth rate of R&D investments with past 1-year returns.

In the regressions, amongst others, I condition on the firm's gross profitability[3,4]. On the one hand, investments in the R&D are associated with higher future economic profits and dividends which are increasing future returns. On the other hand, R&D investments are expensed in the same period as they are incurred which is decreasing earnings (net income before extraordinary items), and hence future returns. The gross profits which are defined as the revenues minus the costs of goods sold do not include R&D expenditures and measure the accounting profitability of R&D intensive firms before R&D investments are taken into account. An appealing feature of gross profitability is that it can be decomposed in R&D-to-assets ratio and the inverse of another measure which is measuring R&D intensity, R&D-to-gross profits.

Chan et al. (2001) try to explain higher future returns of firms with high R&D-to-market value ratios by arguing that managers are investing large R&D outlays despite the firm has performed poorly because they are relatively optimistic about the firm's future prospects. Market participants do not share the management's insight about the firm's business

---

[3] In addition to gross profitability, in the cross sectional regressions I condition on book-to-market, size and 1-month past return.

[4] The gross profitability is a factor in Fama and French's (2015) five-factor pricing model which was uncovered by Novy-Marx (2013).

operations, discount the poor past performance too heavily, and are sluggish in revising their expectations[5].

My empirical findings suggest that when managers are reluctant to cut the *level* of R&D expenditures despite the poor past performance the firm is rewarded with higher subsequent returns. When managers decide to *increase* their R&D expenditures, on the other hand, the good track record in the past price performance is providing the signal for higher future returns. Only firms that have demonstrated their ability to make good investment decisions, and therefore exerted positive price performance over the last year, are rewarded with higher future returns when they increase their R&D expenditures.

Interaction of R&D investments (in levels or changes) with past returns is, however, unable to explain higher future returns of firms with higher R&D-to-market value ratios. Firms with high R&D-to-market value ratios seem to be rewarded with higher future returns which are unrelated to an interaction of R&D investments with past returns.

The rest of the paper is organized as follows. Section 2 reviews the literature on the interaction of R&D investments with past returns. Section 3 presents the data. Section 4 presents my main asset pricing results. Section 5 concludes.

---

[5] In the introduction of their paper Chan et al. (2001) write: "The strongest signs of an association between R&D intensity and future returns come from stocks with high R&D relative to market value of equity (that tend to have experienced poor returns in the past) ... The market apparently gives insufficient credit to past losers who are spending heavily on R&D. Such firms probably face strong pressures to cut R&D and improve earnings. Their reluctance to do so, however, may reflect their managers' confidence that future prospects are not so bleak. Nonetheless, the market tends to overlook such signals (just as it tends to discount other indicators of managers' optimism such as stock repurchases and insider trades)."

2. Literature review of the interaction of R&D investments with past returns

Research has shown that it is important to take into account past performance for quantifying the effect of R&D investments on future stock returns. Chan et al. (2001), for instance, find that firms with high R&D-to-market value ratio on average earn higher future returns. They find that firms in the fifth R&D-to-market value ratio (RDM) portfolio earn an average annual return which is 11.1 percentage points higher than firms in the first RDM portfolio over the first three years after the portfolio formation.

The difference in average annual three-year return shrinks to 1.3 percentage points when firms are sorted based on the R&D-to-sales ratio which does not include the past performance in the denominator. Chan et al. (2001) conclude that RDM has a better predictive power for future returns than R&D-to-sales (RDS).

The stronger predictive power of the RDM relative to the R&D-to-sales ratio could be driven by a past performance which is in the denominator of the RDM. To investigate how past performance interacts with R&D investments, Chan et al. (2001) perform double sorts on R&D-to-sales and past returns. The difference in the future returns is the highest in the portfolio of firms which rank the highest based on the R&D-to-sales ratio and have experienced low vs. high past returns.

The double-sort captures the joint effect of the R&D investments and the past return on the future stock returns. It does not, however, answer the question if the predictive power for future returns is different when *level* or *changes* of R&D investments interact with the past performance, and if the interaction is subsuming the power of R&D-to-market value ratio to predict future returns.

Firms which rank the highest in terms of R&D-to-sales and rank in low past 3-year return portfolio on average earn three-year return which is 4.3 percentage points higher than firms which rank in the high 3-year past return portfolio. The difference shrinks to 0.7 percentage points in the first R&D-to-sales portfolio and is negative in the second R&D portfolio (−0.6 percentage points).

To measure a difference in the future returns of R&D intensive firms which have experienced low vs. high past return, Chan et al. (2001) perform a similar analysis in terms of excess returns. They match the firm with the control portfolio based on size, book-to-market, and past 3-year return. In Table 3 of their paper, Chan et al. (2001) report that the firm's average *excess* 3-year return which ranks in the fifth portfolio based on the R&D-to-sales equals 2.4 percentage points. In Table 5 of their paper, Chan et al. (2001) report that the average excess return increases to 4.4 percentage points when the firm is additionally ranked in the low past 3-year return portfolio. The unexplained difference in returns, therefore, equals 2 percentage points.

The average excess 3-year return of a firm which ranks in the first portfolio based on the R&D-to-sales equals −0.8 percentage points. The average excess return increases to −0.7 percentage points when the firm is additionally ranked in the low past 3-year return portfolio. The difference, therefore, shrinks to 0.1 percentage point. The difference changes direction to −1.5 percentage points in the second R&D-to-sales portfolio (similarly as when I compare the difference in average returns of firms in low minus high past return portfolio which rank in the second R&D-to-sales portfolio).

Although Chan et al.'s (2001) empirical findings clearly suggest that there exists a relationship between the R&D investments and past performance, based on the empirical analysis presented in Chan et al. (2001), one cannot conclude that the average return of a firm with high *level* of R&D expenditures which has experienced poor past performance is statistically significantly different from the average return of the firm which *changes* R&D investments and has experienced lower past performance. Eberhart et al. (2004), for instance, provide empirical evidence showing that changes in R&D-to-assets ratio provide value relevant information when they are significant (when R&D-to-assets ratio changes by more than 5 percent).

In this paper, I focus on the interaction of the level of R&D expenditures which I measure with R&D-to-assets ratio, with past returns, and changes in R&D expenditures which I measure with yearly growth rates with past returns. Past returns could affect differently future returns of firms which have a high level of R&D investments and returns of those firms which have recently increased their R&D spending.

Additionally, I test if the interaction of R&D investments (in levels and changes) with past returns is explaining higher future returns of firms with higher R&D-to-market value ratios.

3. Data

I follow Fama and French (2006) in constructing the book-to-market ratio (BM), market capitalization (MC) and cleaning the sample. B_t, book equity, is total assets (AT) minus liabilities (LT) plus balance sheet items deferred taxes and investment tax credit (TXDITC) if available, minus preferred stock liquidating value (PSTKL) if available, or redemption value (PSTKR) if available, or carrying value (PSTK). Book equity in the regression for July of year $t+1$ is for the fiscal year ending in the calendar year t. The size variable, MC_t, is measured at the end of June of year $t+1$. When I construct B_t/M_t, I measure M_t at the end of December of the year t.

I drop financial firms (Standard Industrial Classification codes between 6000 and 6999) and small firms with total assets less than or equal to $25 million, or book equity less than or equal to $12.5 million in year t to mitigate in sample small firm effect. I exclude firms with negative book equity or price in period t. Additionally, to be included in the sample, a firm must have Compustat data for year t on book equity, revenues, costs of goods sold, R&D expenditures, shares outstanding, and total assets. A firm must have a market cap (price times shares outstanding) available in the Center for Research in Security Prices (CRSP) database for December of t and June of $t+1$.

Finally, the firm must have a dependent variable, monthly return, available in CRSP monthly database. I follow Novy-Marx (2013) and measure the past performance at the short horizon over the last month ($r_{0,-1}$) and over the last 12 to 2 months ($r_{-2,-12}$). Construction of $r_{0,-1}$ and $r_{-2,-12}$ requires that firm in period t has prices for $t, t-1, t-2$ and $t-12$ available in CRSP monthly database. I follow Fama and French (1995) and compute price return as a change in CRSP stock price from period t to $t-\tau$, or RETX variable where available. I drop firms with zero R&D expenditures. I *do not* drop observations of firms with negative R&D (Compustat data item XRD) following Wharton Compustat editors' FAQ 63[6]. Table 1 presents descriptive statistics.

Table 1: Descriptive statistics. Sample spans from December 1975 to June 2014.

Variable	Number of Observations	Mean	Standard Deviation	Percentiles 5th	Percentiles 95th
Market Capitaliazion	414,964	2,044,817	6,858,517	24,023	9,075,230
Book-to-market ratio	414,964	1.2	4.5	0.1	2.3
Gross Profitability	414,964	37.9%	23.0%	4.8%	76.2%
R&D-to-gross profits	414,964	0.5	49.1	0.01	0.6
R&D-to-assets (RDA)	414,964	0.074	0.086	0.004	0.235
R&D-to-market value (RDM)	414,964	0.092	0.308	0.005	0.270
Yearly growth of R&D investments (delta R&D)	338,234	21.6%	98.5%	-30.8%	89.5%
$r_{-2,-12}$	414,964	9.0%	58.1%	-58.2%	100.0%
$r_{0,-1}$	414,964	1.1%	14.7%	-20.9%	25.1%

Table 1 reveals that the average firm in the sample has 2 billion US Dollar of market capitalization and book-to-market ratio which equals 1.2. The average firm has gross

[6] "As explained by the data vendor (also confirmed by our internal investigation), the cases of negative XAD and XRD are very rare. However, these are NOT data errors. For example, if you take a look at the following company (Energy Focus - GVKEY=030614), you will see that it did report two consecutive years of negative R&D expenses. You will need to dig into the actual filing itself to see why this is the case. Most likely, the negative R&D expenses reflect adjustments of funds allocated in the earlier years." Available at https://wrds-web.wharton.upenn.edu/wrds/support/Additional%20Support/WRDS%20Knowledge%20Base%20with%20FAQs.cfm?folder_id=658&article_id=5724

profitability, which is defined as gross profits divided by total assets, which equals 37.9 percent. The standard deviation of gross profitability equals 23 percent.

The R&D-to-assets ratio on average equals 0.074 with a standard deviation of 0.086. The average R&D-to-gross profits ratio equals 0.5 with a standard deviation which equals 49.1. The average R&D-to-market value ratio equals 0.092 with a standard deviation of 0.308. The yearly growth of R&D investments on average equals 21.6 percent with a standard deviation of 98.5 percent. Standard deviations of R&D-to-gross profits ratio and yearly growth rate of R&D investments are higher than standard deviations of R&D-to-assets and R&D-to-market value ratios. I winsorize the distribution of independent variables at 1 and 99 percentiles to mitigate the effect of extreme values (outliers).

Past 1-year return ($r_{-2,-12}$) on average equals 9 percent with a standard deviation of 58.1 percent. Average 1-month return ($r_{0,-1}$) equals 1.1 percent with standard deviation which is equal to 14.7 percent.

4. Main Asset Pricing Results

In the first part of this section, I test which measure of R&D intensity provides better explanatory power for the future returns, (1) R&D-to-market value ratio, (2) R&D-to-assets ratio, (3) R&D-to-gross profits ratio or (4) yearly growth of R&D investments (delta R&D). I present decomposition of gross profitability into R&D-to-assets ratio and R&D-to-gross profits ratio to motivate the gross profitability as a control variable.

In the second part of this section, I use Fama Macbeth cross-sectional regressions to test how R&D investments in levels, measured with R&D-to-assets ratio, and R&D investments in changes, measured with yearly growth rates, interact with past performance. Additionally, I test if the interaction of R&D investments (in levels or changes) with past performance can explain higher future returns of firms with higher R&D-to-market value ratios.

The gross profits-to-assets variable (GPA) is defined in the following way

$$GPA_{i,t} = \frac{Revenues_{i,t} - Costs\ of\ goods\ sold_{i,t}}{Total\ Assets_{i,t}} = \frac{Gross\ profits_{i,t}}{Total\ Assets_{i,t}} \tag{2}$$

Where:

$Revenues_{i,t}$ – Annual revenues of company i in time period t.

$Costs\ of\ goods\ sold_{i,t}$ – Annual costs of goods sold by company i in time period t.

$Total\ Assets_{i,t}$ – Total assets which equal total equity + total liabilities of company i in time period t.

I decompose GPA into R&D-to-assets (RDA) and R&D-to-gross profits (RDGP) in the following way[7]

$$GPA_{i,t} = \frac{RDA_{i,t}}{RDGP_{i,t}} \tag{3}$$

Where:

$$RDA_{i,t} = \frac{R\&D\ Expenditures_{i,t}}{Total\ Assets_{i,t}} \tag{4}$$

$$RDGP_{i,t} = \frac{R\&D\ Expenditures_{i,t}}{Gross\ profits_{i,t}} \tag{5}$$

$R\&D\ Expenditures_{i,t}$ – Annual R&D expenditures as reported in the income statement of company i in time period t.

I define the yearly growth rate of R&D investments (delta R&D) as

$$delta\ R\&D_{i,t} = \frac{R\&D\ Expenditures_{i,t} - R\&D\ Expenditures_{i,t-12}}{R\&D\ Expenditures_{i,t-12}} \tag{6}$$

Since R&D expenditures are available at yearly frequency, the yearly growth rate is equal across 12 months during the year but varies across firms.

Table 2 reports the results of Fama and MacBeth (1973) regressions of monthly stock returns on R&D-to-market value ratio (RDM), R&D-to-assets (RDA), R&D-to-gross profits (RDGP), the yearly growth rate of R&D investments (delta R&D), past 1-year return

[7] In addition to decomposition presented in this paper, the gross profitability can be decomposed as well into an asset turnover and gross margin (Du Pont's model). Novy-Marx (2013) finds that the gross profitability subsumes the power of asset turnover and gross margin to predict future returns.

($r_{-2,-12}$) and gross profitability (GPA). Regressions include controls for size (ln(MC)), book-to-market (BM) and short-horizon past performance ($r_{0,-1}$). The sample spans from December 1975 through June 2014.

Table 2: Fama and MacBeth (1973) regressions of subsequent 1-month stock returns ($r_{1,0}$) on past 1-year return ($r_{-2,-12}$), R&D-to-market value ratio (RDM), R&D-to-assets (RDA), changes in R&D investments (delta R&D), gross profitability (GPA) and R&D-to-gross profits (RDGP). Regressions include controls for the log of market capitalization (ln(ME)), the log of book-to-market (ln(B/M)) and prior month return ($r_{0,-1}$). Independent variables are winsorized at 1% and 99% levels. The sample covers December 1975 through June 2014, the year when FAS No. 2 (1974) came into effect. t-statistics are presented in the second rows. Slope coefficients with t-statistics above 1.96 (in absolute terms) are bolded. Regressions use Newey-West standard errors (lag 12).

Average R2	5.7%	5.3%	5.3%	5.5%	5.9%	6.0%	6.4%
Number of time periods	451	451	439	451	451	439	439
Average number of firms	920	920	770	920	920	770	770
Independent variable	Estimated coefficients (t-values in second rows) (1)	Estimated coefficients (t-values in second rows) (2)	Estimated coefficients (t-values in second rows) (3)	Estimated coefficients (t-values in second rows) (4)	Estimated coefficients (t-values in second rows) (5)	Estimated coefficients (t-values in second rows) (6)	Estimated coefficients (t-values in second rows) (7)
R&D-to-assets (RDA)	**0.0338**				0.0172		0.0145
	2.25				1.01		0.93
R&D-to-gross profits (RDGP)		0.0080					
		1.60					
Delta R&D			-0.0015			**-0.0019**	**-0.0020**
			-1.61			-2.18	-2.33
R&D-to-market value (RDM)				**0.0023**	**0.0018**	**0.0020**	**0.0015**
				2.89	3.60	2.46	2.91
GPA	**0.0092**	**0.0092**	**0.0073**	**0.0079**	**0.0082**	0.0059	**0.0065**
	3.07	3.06	2.48	2.51	2.66	1.86	2.05
$r_{-2,-12}$	0.0036	0.0038	0.0023	0.0038	0.0037	0.0017	0.0016
	1.45	1.53	0.87	1.55	1.52	0.66	0.61
log(ME)	**-0.0011**	**-0.0014**	**-0.0014**	**-0.0011**	**-0.0011**	-0.0011	-0.0010
	-2.15	-2.45	-2.21	-2.10	-2.14	-1.95	-1.94
log(BM)	**0.0030**	**0.0023**	0.0012	0.0011	0.0017	0.0001	0.0007
	3.22	2.20	1.12	0.85	1.69	0.04	0.66
$r_{0,-1}$	**-0.0475**	**-0.0462**	**-0.0476**	**-0.0483**	**-0.0489**	**-0.0489**	**-0.0494**
	-7.97	-7.44	-7.38	-7.93	-8.14	-7.82	-8.05

Specification 1 in Table 2 shows that the R&D-to-assets ratio has a positive average effect on the future 1-month return which is statistically significant. In specifications 2 and 3 we can see that the average cross-sectional slopes of R&D-to-gross-profits and changes in R&D investments are not significantly different from zero. In specification 4, we can see that the average effect of R&D-to-market value is positive and statistically significant. These empirical findings suggest that the R&D-to-assets and R&D-to-market value ratios have positive predictive power for future stock returns. Changes in R&D investments and R&D-to-gross profits, on the other hand, do not have predictive power for future returns.

I have estimated the cross-sectional regression which includes the R&D-to-sales and found, similarly as with the R&D-to-gross profits, that it has an average effect on the 1-month future return which is insignificant[8]. Difference between R&D-to-gross profits ratio and R&D-to-sales ratio is that in case of the R&D-to-gross profits costs of goods sold are subtracted from the revenues in the denominator.

I follow Novy-Marx (2013), Chan et al. (2001), and Chambers et al. (2002) and do not condition on the firm's beta. Lev and Sougiannis (1996) estimate the cross-sectional regressions which condition on the firm's beta and show that the coefficient on the R&D-to-market value ratio remains positive and statistically significant.

Specification 5 in Table 2 compares the predictive power of R&D-to-assets ratio and R&D-to-market value. R&D-to-market value is subsuming the predictive power of R&D-to-assets ratio. The average cross-sectional slope of R&D-to-assets ratio is halved and becomes insignificant. Therefore, R&D-to-assets ratio is not explaining the positive predictive power of R&D-to-market value for future stock returns.

To further investigate the interaction of R&D investments with past performance, specification 6 tests the predictive power of *changes* in R&D investments if I condition on

[8] In the regression which is similar to regression in specification 2 in Table 2 but includes R&D-to-sales instead of R&D-to-gross profits, the estimated average coefficient on the R&D-to-sales equals 0.0062 with t-statistic equal to 0.61.

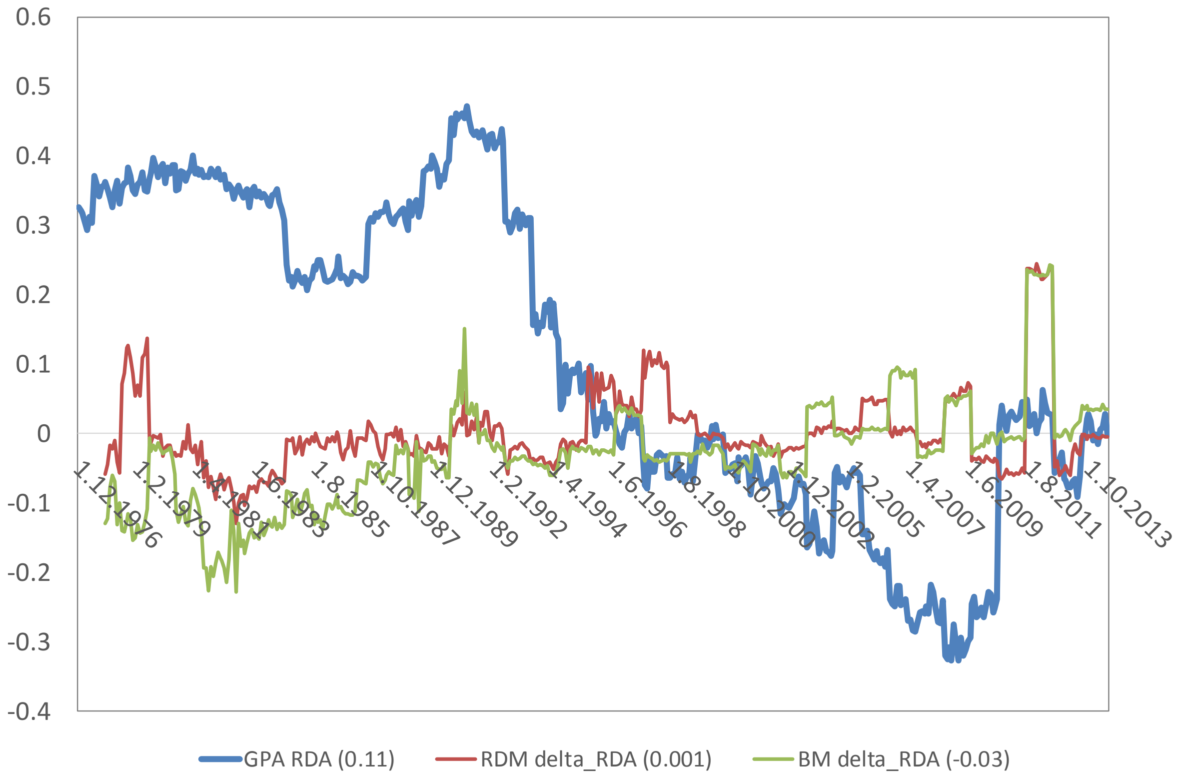
R&D-to-market value ratio. We can see that the negative predictability of changes in R&D investments strengthens and becomes statistically significant when I condition on the R&D-to-market value ratio. Specification 7 shows that average slopes of changes in R&D investments and R&D-to-market value ratio do not change when I additionally condition on the R&D-to-assets ratio.

Therefore, these empirical findings suggest that higher R&D-to-market value predicts higher future returns which are unrelated to R&D-to-assets ratio or changes in R&D investments. While firms with high R&D-to-market value ratio on average earn higher returns, conditioning on the R&D-to-market value decreases future returns of firms which change R&D investments. The firm is rewarded with higher future returns if it has high R&D-to-market value ratio while my empirical findings suggest that the firm is on average penalized with lower future returns if it increases R&D investments.

Before I continue by interacting changes in R&D investments and R&D-to-assets ratio with past returns, Figure 1 depicts the cross-sectional correlations between gross profitability (GPA), R&D-to-assets (RDA), changes in R&D investments (delta R&D), R&D-to-market value ratio (RDM), and book-to-market ratio (BM). Correlation between gross profitability and R&D-to-assets is on average positive (0.11) but has in the year 1996 switched from being positive to negative. Therefore, while before 1996 firms which had higher R&D-to-assets were associated with higher gross profitability, after 1996 such an association has vanished.

While average slope on changes in R&D investments becomes significant if I condition on R&D-to-market value ratio (specification 6 in Table 2), Figure 1 shows that changes in R&D investments and R&D-to-market value ratio are essentially uncorrelated. Time series average of cross-sectional correlations between changes in R&D investments and R&D-to-market value ratio equals 0.001 (red curve in Figure 1). The average cross-sectional correlation between the book-to-market ratio and changes in R&D investments is slightly higher in absolute terms and on average equals -0.03 (green curve in Figure 1). It exhibits similar time-series pattern as the cross-sectional correlation between changes in R&D investments and RDM ratio.

Figure 1: Cross-sectional correlations between gross profitability (GPA), R&D-to-assets (RDA), R&D-to-market value ratio (RDM), and changes in R&D investments (delta R&D). Time averages of cross-sectional correlations are presented in brackets.



In order to test the full effect of past returns, R&D investments and their potential mutual reinforcement on the future returns, I extend the model with an interaction term between R&D investments and past returns. Table 3 presents the results of cross-sectional regressions. I follow Novy-Marx (2013) and measure (long-run) past performance with the past 1-year return. In the cross-sectional regressions, I have tested the predictive power of the past 3-year return and found that in the sample of firms which report R&D expenditures it has lower predictive power for the future returns than the past 1-year return[9]. I should stress out that in the cross-sectional regressions, I condition on the short-run past performance, $r_{0,-1}$, which is highly statistically significant, and takes the major bulk of the return reversal out of the future 1-month return.

[9] In the regression which is similar to regression in specification 1 in Table 3 but includes the past 3-year performance instead of the past 1-year performance, the estimated average coefficient on the past 3-year performance equals -0.0011 with t-statistic equal to -0.71.

Table 3: Fama and MacBeth (1973) regressions of subsequent 1-month stock returns ($r_{1,0}$) on past 1-year return ($r_{-2,-12}$), R&D-to-assets (RDA), R&D-to-market value ratio (RDM), changes in R&D investments (delta R&D), the interaction term between R&D-to-assets and past returns (R&D-to-assets $\times r_{-2,-12}$), and the interaction term between changes in R&D investments and past returns (delta R&D $\times r_{-2,-12}$). Regressions include controls for the gross profitability (GPA), log of market capitalization (ln(ME)), the log of book-to-market (ln(B/M)) and prior month return ($r_{0,-1}$). Independent variables are winsorized at 1% and 99% levels. The sample covers December 1975 through June 2014, the year when FAS No. 2 (1974) came into effect. t-statistics are presented in the second rows. Slope coefficients with t-statistics above 1.96 (in absolute terms) are bolded. Regressions use Newey-West standard errors (lag 12).

Average R2	4.9%	6.2%	5.4%	6.2%	5.4%
Number of time periods	451	451	451	439	439
Average number of firms	920	920	920	770	770
Independent variable	Estimated coefficient (p-values in parenthesis) (1)	Estimated coefficient (p-values in parenthesis) (2)	Estimated coefficient (p-values in parenthesis) (3)	Estimated coefficient (p-values in parenthesis) (4)	Estimated coefficient (p-values in parenthesis) (5)
$r_{-2,-12}$	0.0040	**0.0057**	**0.0058**	0.0002	0.0005
	1.61	2.01	2.17	0.07	0.21
R&D-to-assets (RDA)		0.0234	0.0206		
		1.31	1.20		
RDA $\times$ $r_{-2,-12}$		**-0.0499**	**-0.0517**		
		-2.37	-2.52		
Delta R&D				**-0.0024**	**-0.0024**
				-2.61	-2.69
Delta R&D $\times$ $r_{-2,-12}$				**0.0094**	**0.0086**
				4.42	4.22
R&D-to-market value (RDM)		**0.0018**	**0.0019**	**0.0021**	**0.0021**
		3.59	3.76	2.49	2.50
GPA	**0.0089**	**0.0082**	**0.0078**	0.0060	0.0055
	2.96	2.66	2.61	1.88	1.75
log(ME)	**-0.0014**	**-0.0011**	**-0.0012**	**-0.0011**	**-0.0012**
	-2.42	-2.14	-2.39	-1.97	-2.19
log(BM)	**0.0022**	0.0016	0.0016	0.0000	0.0001
	2.07	1.70	1.78	0.03	0.06
$r_{0,-1}$	**-0.0458**	**-0.0489**		**-0.0491**	
	-7.42	-8.17		-7.78	

Specification 1 in Table 3 shows that past 1-year performance does not have a statistically significant average effect on the future 1-month return once I condition on the past 1-month return. The average estimated slope of the short-run performance ($r_{0,-1}$) is *negative* and statistically significant. Therefore, a bulk of return reversal comes from the previous month's return.

In specification 2 in Table 3, I test the predictive power of the interaction term between the RDA and the past 1-year performance on the future 1-month return. I define the interaction term as the R&D-to-assets ratio which I multiply by the past 1-year return:

$$Interaction\ term\ RDA_{i,t} = RDA_{i,t} \times r_{-2,-12,i,t} \tag{7}$$

We can see that the average slope of the interaction term between R&D-to-assets ratio and past 1-year performance is negative (specification 2). Therefore, these empirical findings support the argument that firms which have high R&D-to-assets ratio and have experienced poor past performance on average earn higher future returns. In specification 3, we can also see that the average slope changes marginally if I do not condition on the past 1-month return (which is in the last row of Table 3).

Last two specifications in Table 3 (specification 4 and 5) test the predictive power when changes in R&D investments are interacted with past 1-year returns:

$$Interaction\ term\ \Delta\ R\&D_{i,t} = \Delta\ R\&D_{i,t} \times r_{-2,-12,i,t} \tag{8}$$

In specification 4, we can see that changes in R&D investments preserve their negative predictive power for future returns. The firm which increases R&D investments and does not experience higher past returns earns lower future returns (average slope equals -0.0024). Interestingly, the interaction term between changes in R&D investments and past 1-year performance has positive predictive power for future stock returns (average slope equals

0.0094). Specification 4 shows that firms which increase R&D investments are rewarded with higher future returns only if they have experienced higher past 1-year returns.

Specification 5 in Table 3 shows that the average cross-sectional slope of changes in R&D investments remains unchanged if I do not condition on the past 1-month performance. Similarly, the average cross-sectional slope of the interaction term between changes in R&D investments and past 1-year performance is marginally changed when I do not condition on the past 1-month performance. It decreases from 0.0094 to 0.0086.

I have checked the robustness of results presented in specifications 2 and 4 in Table 3 using the weighted least squares estimator. I weight the contribution of each firm in the cross-section by its size, log of its market value. In specification 2, the estimated average cross-sectional slope of the interaction term marginally decreases, to −0.0513. In specification 4, the estimated average cross-sectional slope of the interaction term marginally increases, to 0.0096. Hence, the estimated average effects of the interaction terms do not change significantly when the observations are weighted by the firm's size.

My empirical findings suggest that the interaction of R&D investments with past performance is important for future stock returns. Firms with high R&D investments relative to their assets (high R&D-to-assets ratio) which have experienced poor past performance (negative 1-year past returns) earn higher future stock returns. On the other hand, firms which increase their R&D investments are rewarded with higher future returns only if they have experienced higher past 1-year returns. The information in past 1-year performance, therefore, seems to provide value relevant signal for future returns of firms with a high level of R&D investments and firms which are changing (increasing) their R&D investments.

Specifications 2 to 5 also show that the interaction of R&D investments (in levels or changes) with past performance is not explaining higher returns of firms with high R&D-to-market value ratios. R&D anomaly, namely that firms with high R&D-to-market value ratios earn higher returns remains unexplained after interacting R&D investments with past performance.

5. Conclusions

In this paper, I investigate how the interactions of R&D investments (in levels and changes) with past returns affect future stock returns. I find that firms with a high level of R&D investments (high R&D-to-assets ratios) earn higher future returns if they have experienced lower past returns. Firms which increase R&D investments, on the other hand, are rewarded with higher future returns only if they have experienced higher past returns.

When managers are reluctant to cut the level of R&D expenditures despite the poor past performance the firm is rewarded with higher subsequent returns. On the other hand, when managers increase the R&D outlays the good track record in the past price performance is providing the signal for future returns. Only firms which have demonstrated their ability to make good investment decisions, and therefore earned positive returns over the last year are rewarded with higher future returns.

Interaction of R&D investments (in levels or changes) with past returns is, however, unable to explain higher future returns of firms with high R&D-to-market value ratios. Higher returns of firms with high R&D-to-market value ratios are unrelated to the interaction of R&D investments with past returns. I am unable to provide empirical support for the argument that higher future returns of firms with high R&D-to-market value ratios are related to firms with a high level of R&D investments which have experienced lower past returns.

Past performance is providing value-relevant signals for future returns of firms with either higher level of R&D investments or higher growth rates of R&D spending. Higher returns of firms with high R&D-to-market value ratios, however, remain unexplained after taking into account the interaction of R&D investments with past returns.

References

1. Aboody, D., & Lev, B. (2000). Information asymmetry, R&D, and insider gains. *The journal of Finance*, 55(6), 2747-2766.
2. Chambers, D., Jennings, R., & Thompson, R. (2002). Excess returns to R&D–intensive firms. *Review of Accounting Studies*, *7*(2-3), 133–158.
3. Chan, L., Lakonishok J., & Sougiannis, T. (2001). The Stock Market Valuation of Research and Development Expenditures. *Journal of Finance*, *56*(2), 431–57.
4. Eberhart, A. C., Maxwell, W. F., & Siddique, A. R. (2004). An examination of long-term abnormal stock returns and operating performance following R&D increases. *The Journal of Finance*, 59(2), 623-650.
5. Fama, E. F., & French, K. R. (1995). Size and book-to-market factors in earnings and returns. *The journal of finance*, 50(1), 131-155.
6. Fama, E. F., & French, K. R. (2006). Profitability, Investment and Average Returns. *Journal of Financial Economics*, 82(3), 491–518.
7. Fama, E. F., & French, K. R. (2015). A Five-Factor Asset Pricing Model. *Journal of Financial Economics*, 116(1), 1–22.
8. Fama, E. F., & MacBeth, J. (1973). Risk, Return and Equilibrium: Empirical Tests. *Journal of Political Economy*, *81*(3), 607–636.
9. FASB (Financial Accounting Standards Board). (1974). Accounting for Research and Development Costs. Statement of Financial Accounting Standards No.2.
10. Lev, B., & Sougiannis, T. (1996). The Capitalization, Amortization and Value–Relevance of R&D. *Journal of Accounting and Economics*, *21*(1), 107–38.
11. Lev, B., & Sougiannis, T. (1999). Penetrating the Book–to–Market Black Box: The R&D Effect. *Journal of Business Finance & Accounting*, *26*(3–4), 419–449.
12. Novy-Marx, R. (2013). The Other Side of Value: The Gross Profitability Premium. *Journal of Financial Economics*, 108(1), 1–28.
13. Sougiannis, T. (1994). The Accounting Based Valuation of Corporate R&D. *The Accounting Review,* 69(1), 44–68.